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The Faculty of Economics, Business Administration and Information Technology of the University of Zurich hereby authorises the printing of this Doctoral Thesis, without thereby giving any opinion on the views contained therein.

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Chairman of the Doctoral Committee: Prof. Dr. Dieter Pfaff

Für meine Eltern

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Preface

Throughout my work on this thesis, I have come across many remarkable individuals which have supported me directly and indirectly. I am finally in the enjoyable position to fulfill the due obligation to express my gratitude.

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Zürich, August 2010 Hannes Ullrich

Introduction

This dissertation is a collection of three essays on topics in applied microeconometrics, transport economics, and applied microeconomic theory.

At least since Hannibal's invasion into Italy, the difficulties of crossing the Alps have been revealed. Overcoming this natural bottleneck requires traveling across high-altitude passes or, somewhat more recently, through road and rail tunnels that are expensive to build and maintain. Even today, safety remains a vital political issue fueled by repeated fatal accidents at various alpine crossings. The first part of this thesis sheds light on the economics behind freight route choice and investment in road and rail infrastructure. Chapters 1 and 2 evaluate hypothetical changes in the availability of transalpine transport infrastructure in terms of substitution patterns and effects on consumer welfare.

In Chapter 1, we analyze the expected effects of building a transalpine rail tunnel between Lyon and Turin on i) the market shares of incumbent and entrant freight transport service suppliers, and ii) consumer surplus. Such a large infrastructure project, the Lyon-Turin Transalpine, has been planned for decades and preparatory construction works have begun recently. It consists of a 53km rail tunnel providing freight shippers with a new alpine path. Modeling transport decisions such as route choice and pricing is a complex task. Data including all complexities are rare and a model's virtue is to reduce complex reality to economically relevant trade-offs. We aim to shed light on inter-modal transalpine freight transport decisions having merely aggregate volume measures at hand. We employ a simple numerical equilibrium model following Ivaldi and Vibes (2008), where freight shippers choose a mode and alpine path to ship goods from a given origin to a given destination. Freight carriers strategically set prices for the differentiated products they supply. The base decision model is a nested logit discrete choice model as in, for example, McFadden (1973), Berry (1994), and Ortúzar (2001). The nested logit model has a closed-form solution for individuals' choice probabilities which, in markets with many consumers, can be approximated by aggregate market shares. Due to its analytical simplicity, we can calibrate this demand model and the supply side to observed market shares, prices, route characteristics, and cost measures. Using the equilibrium parameters, demand elasticities and within-nest correlation, we perform counterfactual simulations to investigate demand and supply side reactions to the introduction of a new rail alternative.

For our analysis, we define the following three geographical sub-markets: a regional short-distance market (Lyon - Turin, 315km), a wider North-South market (Paris - Milano, 850km), and a West-East market (Madrid - Milano, 1575km). We find limited substitutability between freight transport products on a North-South transit axis. On the Paris-Milano market, the shippers' decision remains largely based on mode choice. The new high-quality rail alternative attracts new demand but does not succeed in lowering road demand. When we shorten this axis to the regional market between Lyon and Turin, both a modal shift and an increase in demand for shipping occur, showing the same variations in the market share of the outside option and consumer surplus as in the Paris-Milano market. In contrast to the North-South axis, the West-East transit market appears a better candidate for modal shift. Between Madrid and Milano, the new rail link appears sufficiently attractive for shippers to switch modes. Overall traffic does not increase after the introduction of the new link, suggesting higher volatility of shippers' preferences across products on this transit axis. Should European rail integration be fostered, the new transalpine link between Lyon and Turin could play a complementary part among other projects. It would be of interest to compare the respective impacts on the West-East transit axis of the Lyon-Turin Transalpine and the Perpignan-Figueras Transpyrenees between France and Spain.

Based on our analysis, the construction of a new high quality infrastructure may only be one tool out of a global modal shift-oriented policy toolbox. For the French-Italian alpine corridor, more direct and committed intervention based on a variety of policy measures may open a fruitful path to the political goal of increasing modal shift towards rail. We show that a scheme as observed in Switzerland, where cross-subsidies from rail to road generate incentives for modal shift, are a valid complement to investment in new rail infrastructure and even, to some extent, a reasonable substitute.

In Chapter 2, we further investigate consumer welfare in transalpine freight transport using micro-data on individual route choice. We tackle two core questions in this chapter. First, to what extent does the way we model unobserved heterogeneity matter for welfare estimates in discrete choice models? Second, what is the loss in consumer surplus per year from shutting down a transalpine road infrastructure such as the Mont Blanc tunnel? Closing this tunnel has been increasingly proposed in the political debate following several fatal road accidents in large alpine tunnels.

The most severe accident, in the Mont Blanc in 1999, led to a full closure of the tunnel over a period of 3 years. As in Chapter 1, we model route choice as a discrete choice among a number of mutually exclusive alternatives. Due to our rich data, we can flexibly model unobserved heterogeneity of decision makers in their valuation of money and time. Decision makers may be heterogenous for a number of reasons. For example, the value of alternatives may depend on the weight or type of commodity a truck is transporting. These are examples for observable heterogeneity which are easily controlled for. However, the value of money and time is also likely to depend on unobservable truck characteristics. These could be *en route* pick-ups of goods, special logistic needs, or truck drivers' personal tastes that favor one route over another. Modeling such unobserved heterogeneity in the discrete choice framework has been at the heart of research analyzing economic choices during the last two decades. The workhorse model has been the random coefficients, or mixed, logit model, presented in detail by McFadden and Train (2000). More recently, researchers have started asking how the way we model unobserved heterogeneity affects policy-relevant measures such as price elasticities or consumer welfare. Hensher and Greene (2003) as well as Cherchi and Polak (2005) are early contributions highlighting the potential importance of these unresolved modeling issues. To our knowledge, the only existing contribution evaluating implications on consumer welfare are Hynes et al. (2008). In a destination choice setting of whitewater sites in Ireland, they analyze differences between the random coefficient model and a latent class model in consumer surplus estimates.

We contribute to this literature by applying a recently proposed flexible nonparametric estimator of unobserved heterogeneity to a random coefficients logit model and investigating the impact of parametric assumptions on a measure of consumer welfare. By definition, we expect the assumptions we make on the distribution of preferences to have an impact on our measure of consumer surplus. Bajari et al. (2007, 2010), henceforth BFKR, propose an estimator of unobserved heterogeneity in a general class of economic models, using sieve methods. In a nutshell, their idea is to approximate the true underlying taste distribution by a finite grid in the preference space. In general, such methods require large sample sizes as the data must supply the model structure. Their motivating example is a random coefficients logit model with individual-level data and they provide convincing Monte Carlo evidence. To our knowledge, we are the first to apply their estimator to real-world data in a static discrete choice model with random coefficients.

To identify the underlying structural parameters, we use a large scale individual choice data set in a transport context. In particular, we use the 2004 Cross-Alpine Freight Transport survey data to investigate the economic choice of transalpine freight traffic. We exploit exogenous variation in travel cost and time arising from the fixed geographic locations of origin, destination, and alpine crossing points. While endogeneity concerns are less important with individual-level data, we discuss several potential sources of endogeneity bias such as congestion or weather conditions. We find that parametric assumptions and the dimensionality of modeled unobserved heterogeneity have a significant impact on welfare results. Our BFKR estimates predict economically significantly higher annual losses in user surplus due to the Mont-Blanc tunnel closure. The latter implies a loss of €5.39 Mio and the parametric random coefficients logit model a loss of €2.97 Mio in specifications where both price and time are assumed to have random coefficients. With one random coefficient, the BFKR estimate is almost double that of the parametric random coefficients logit estimate, €7.09 Mio versus €3.62 Mio. Compared to the logit with fixed coefficients and the BFKR estimates, both parametric random coefficient specifications underestimate the loss in consumer surplus.

In Chapter 3, we contribute to the empirical literature on effort incentives in contests. In many situations, applicants compete for a limited number of positions, and selection is based on *perceived* skill or talent, for example in hiring and promotion procedures or in nominations of election candidates by political parties.

We provide a theory of agents' effort incentives in such situations and test the predictions of this theory using data from professional soccer. Our theory introduces signal jamming as in career concerns models à la Holmström (1982) in rank-order tournaments[1], allowing for asymmetries between agents. We show that incentives are strongest in close contests, i.e., when several agents have similar ex ante winning probabilities. In some contexts, higher effort also increases the risk of an injury and/or leads to exhaustion. When competing for a position that requires continued fitness, candidates who are confident that their reputation sufficiently exceeds that of other contestants may find it optimal to exert less than normal effort.

We test these predictions using a panel data set on the German Soccer League in the seasons 2006/07 and 2007/08. A subset of players belong to nations that qualified for the Euro Cup in summer 2008, the *Euro 2008*, the most prestigious international soccer Cup alongside the World Cup, and thus participated in a *nomination contest*. All other players belong to nations either outside of Europe, thus being excluded from the Euro Cup by definition, or in Europe that did not qualify for *Euro 2008*. The latter serve as a natural control group as they work in exactly the same club environment while facing comparable mechanisms of reputation formation. We thus observe variation that lends very well to a difference-in-difference analysis. Most importantly, we are able to measure reputation asymmetries which, as we will show, enables us to identify incentive effects depending on relative reputation along the entire range of reputations. Our measure is based on participation in past national team games, that is, the number of game nominations by national coaches. We provide evidence that our measure is a good proxy of players' reputations.

[1] See Lazear and Rosen (1981) for the theoretical framework of rank-order tournaments, Ehrenberg and Bognanno (1990) for early empirical evidence, and Prendergast (1999) for a survey.

We are aware of two empirical contributions showing the relevance of player asymmetries on incentive effects in tournaments. They are both examples from golf tournaments. Brown (2010) analyzes performances in tournaments in which players do and do not compete against a real golf superstar, Tiger Woods. She finds that players underperform in tournaments in which they compete against Tiger Woods and, thus, have a close to zero chance of winning. This measure is a binary measure having either a positive value smaller than or equal to one or being equal to zero. In the second paper, Franke (2010) analyzes biased golf tournaments. There are two types of amateur golf tournaments in Germany, gross and net tournaments. In the latter, players are ranked by the deviation from their pre-tournament handicap while in gross tournaments rankings are as in standard professional golf tournaments. He uses this variation to identify performance effects due to varying winning probabilities. Using our continuous proxy of winning probabilities, we are able to track the equilibrium relation between asymmetric reputations and effort more closely. In particular, in contrast to the existing literature, our measure covers the area close and up to probabilities equal to one in a setting where winning the tournament leads to subsequent career opportunities. The latter allows us to identify the effect of injury (or exhaustion) concerns predicted by our theoretical model.

We find a large positive effect of nomination contest participation on several output measures, for example the number of ball contacts, for players with intermediate chances of being nominated. For players whose nominations chances are very high, however, the effect of contest participation is negative. That means that players whose uncertainty over their (non-)nomination is highest will exert the most effort in order to positively influence their national team coaches' nomination decisions. Players who are certain of (not) being nominated do not have any incentive to exert extra effort since it will have no impact on the decision. Much rather do (almost) certainly nominated players reduce effort in club games in order to avoid injuries that may jeopardize their Euro Cup participation.

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Chapter 1

Entry and Competition in Freight Transport: The Case of a Prospective Transalpine Rail Link between France and Italy

joint with Delphine Prady

1 Introduction

The Alps have long posed challenges to European transport infrastructure planners. Overcoming this bottleneck requires high-altitude Alpine passes or road and rail tunnels which are difficult and expensive to build and maintain. A comparably large infrastructure project outside of the Alps has been the Channel Tunnel, for example, connecting France and the United Kingdom via rail. Kay et al. (1989) predict the social and private profitability of the Channel Tunnel project. In this paper, we analyze partial economic returns of public investment in a specific and much debated transalpine rail infrastructure. To do so, we employ a simple equilibrium model and, by numerical simulation, compute the changes in direct infrastructure users' surplus. This approach can ultimately be integrated into a more elaborate cost-benefit analysis.[1]

[1] De Jong et al. (2005) present a survey on welfare evaluation in the discrete choice random utility framework.

More precisely, we analyze the competitiveness of freight transport supply by rail and road carriers on the Lyon-Turin corridor. We perform an equilibrium analysis in the context of a discrete choice model (see Anderson et al., 1992, or McFadden, 1981) which allows to analyze demand and competitive supply of differentiated products. We predict the reactions of all competitors given strategies and consumer behavior when facing a new product. In our case, the new product is a prospective high speed rail link between Lyon and Turin, "La Liaison Ferroviaire Lyon-Turin".[2]

We simulate the entry of this new transalpine link on three different markets: regional transport between Lyon and Turin, transit between Spain and Lombardia, and transit between Ile-de-France-Nord and Lombardia. As an alternative policy measure we simulate a deliberate change in some alternatives' cost structure, that is the introduction of a higher road tunnel fee combined with a hypothetical reduction of rail costs. We find limited substitutability between freight transport products – hereafter defined as a "mode+alpine path" bundle – on the North-South transit axis where shippers' choices remain largely mode-driven. The new high-quality rail alternative does attract new demand but does not succeed in lowering demand for road transport. On the regional market between Lyon and Turin, both a modal shift and an overall increase in demand for shipping occur. The West-East transit market appears the best candidate for modal-shift as, between Spain and the region of Lombardia, the new rail link appears attractive enough for shippers to switch modes. Should European rail integration be fostered, the new transalpine link between Lyon and Turin could play a complementary part among other European projects.

Moreover, based on our analysis, the construction of a new high quality infrastructure may only be one tool out of a global modal shift-oriented policy toolbox. For the French Alpine corridor, more direct and committed intervention based on a variety of policy measures as observed in Switzerland may open a more fruitful path to the political goal of increasing modal shift towards rail.

The paper is organized as follows. In section 2 we describe our demand-and-supply model. We present our data set in section 4 and provide our empirical analysis and simulation results in section 3. Section 6 concludes.

2 Modeling Lyon-Turin Freight Transport

Our goal is to evaluate the prospective changes in social welfare of a rail infrastructure project. In our setting, based on Ivaldi and Vibes (2008), consumers, hereafter called *shippers*,[3] choose a transport mode, that is rail or road, and an alpine path

[2] The gains in speed for freight trains are mainly driven by significantly reduced slopes, namely down to 12% from currently 30%, in the prospective base tunnel. Detailed information is available on the web sites http://www.ltf-sas.com and http://www.transalpine.com/

[3] "Consumers" can also be seen as logistic intermediaries acting on behalf of goods producers.

to carry their goods between two specific regions. Suppliers, hereafter called *freight carriers*,[4] are assumed to compete in prices. We then derive the market equilibrium and provide results of counterfactual experiments.

As a first remark, we concentrate on non-combined[5] transport modes in the present analysis. Indeed, we have learned from several phone interviews with freight logistic firms[6] that such a mode is of little importance for most of their shipping activity. To a large extent, that is a peculiarity to the French transport sector in which, historically, the road has been the dominant mode of freight transport. Note also that freight services are not homogenous goods but consist of a widely diversified set of goods with specific haulage requirements and logistic needs. This heterogeneity is to some extent accounted for by equilibrium prices set according to commodity characteristics (such as freshness or hazardousness) which make one choice alternative more attractive than another. Given that we have information on broad commodity classes and aggregate price data only, we conduct our analysis based on these aggregate data. Our model is, in this respect, stylized and simplified but would lend to a more detailed analysis if more disaggregate data were available. In practice, transaction price data on transport services are particularly hard to obtain on a broad basis. Despite this limitation, aggregate data do contain information allowing to contribute policy-relevant conclusions.

Furthermore, as a spatial concern, one needs to distinguish between transit and regional freight transport. Transit and short-distance freight carriers exhibit different company characteristics. For example, long-haul transit freight carriers are rather firms with more than 50 employees, short- and mid-haul freight carriers are rather smaller firms.[7] Accounting for distance-differentiated markets enables us to better characterize competition between products. Indeed, it is unclear whether inter-modal competition is fiercer on long or short distance freight transport markets. Hence, we consider three types of markets targeted by suppliers that can be subsumed into two broader market categories, transit and regional transport. “Transit North-South” (freight traffic between Ile-de-France-Nord[8] and Lombardia) and “Transit East-West” between Spain and Lombardia were chosen based on their important relative shares of French-Italian transalpine passages. We define as short-distance transport freight traffic between the area of Lyon and the area of Turin. Freight traffic on these three markets amounts to 12.6% of total alpine freight traffic reported during the studied period. While this appears a low figure, the traffic scale

[4]For example, SNCF and Trenitalia for the Lyon-Turin rail project.

[5]Combined modes can be both accompanied and unaccompanied transport, where the former corresponds to piggyback transport and the latter to intermodal container transports. Piggyback transport is non-existent in our sample.

[6]GEODIS Calberson GE, GEFCO Network, among others.

[7]London Economics (2003)

[8]Ile-de-France-Nord defines the metropolitan area of Paris and neighboring regions but excludes the Benelux countries.

should not affect competition between the different alpine products. Rather than traffic volume, geographic coverage matters in a product's competitiveness. Figure 1 illustrates these three markets.

2.1 Demand side

Assume a shipper takes a two-step decision:

- first, she decides which mode she wants to carry her commodity with;
- second, she chooses an alpine path.[9]

The second step is motivated as follows. A shipper has an *a priori* ranking of paths. This idiosyncratic ranking is based on characteristics of the shipped goods and of alpine paths. Different paths exhibit different technical characteristics, as detailed below. An alternative is thus a combination of a transport mode and a path to cross the Alps. Product differentiation is mainly due to geographical and regulatory aspects.[10] We assume that shippers have in mind these qualitative differences of available products when sending their goods on a given origin-destination journey (hereafter O–D). In addition to competing differentiated mode-path combinations, we assume the existence of an outside good, OG. It accounts for shippers that are interested in transporting their goods across the Alps by rail or road, yet currently do not. Thus, it represents a potential niche suppliers can target, inducing additional traffic. This is important to bear in mind since aggregate consumer welfare increases with additional traffic.

We classify J alternatives into G groups, where $g = 0, 1, 2$. Group 0 corresponds to the OG, 1 and 2 correspond to rail and road, respectively. Shipper i's utility associated with alternative j is:

$$U_{ij} = V_j + \epsilon_{ij} \tag{1}$$

where:

- V_j is the mean utility level common to all shippers,
- ϵ_{ij} corresponds to the departure of shipper i from the common utility level (also called random part, that is shipper i's unknown idiosyncratic taste for product j).

The random component leading to a nested logit demand model can be further decomposed as follows:

$$\epsilon_{ij} = \sigma\, \nu_{ig} + (1 - \sigma)\, \nu_{ij} \tag{2}$$

[9]The available alpine path are: Mont Blanc, Frejus, Vintimille, Montgenevre, or Mont-Cenis.

[10]Specific haulage requirements, logistic needs, and conveyed goods.

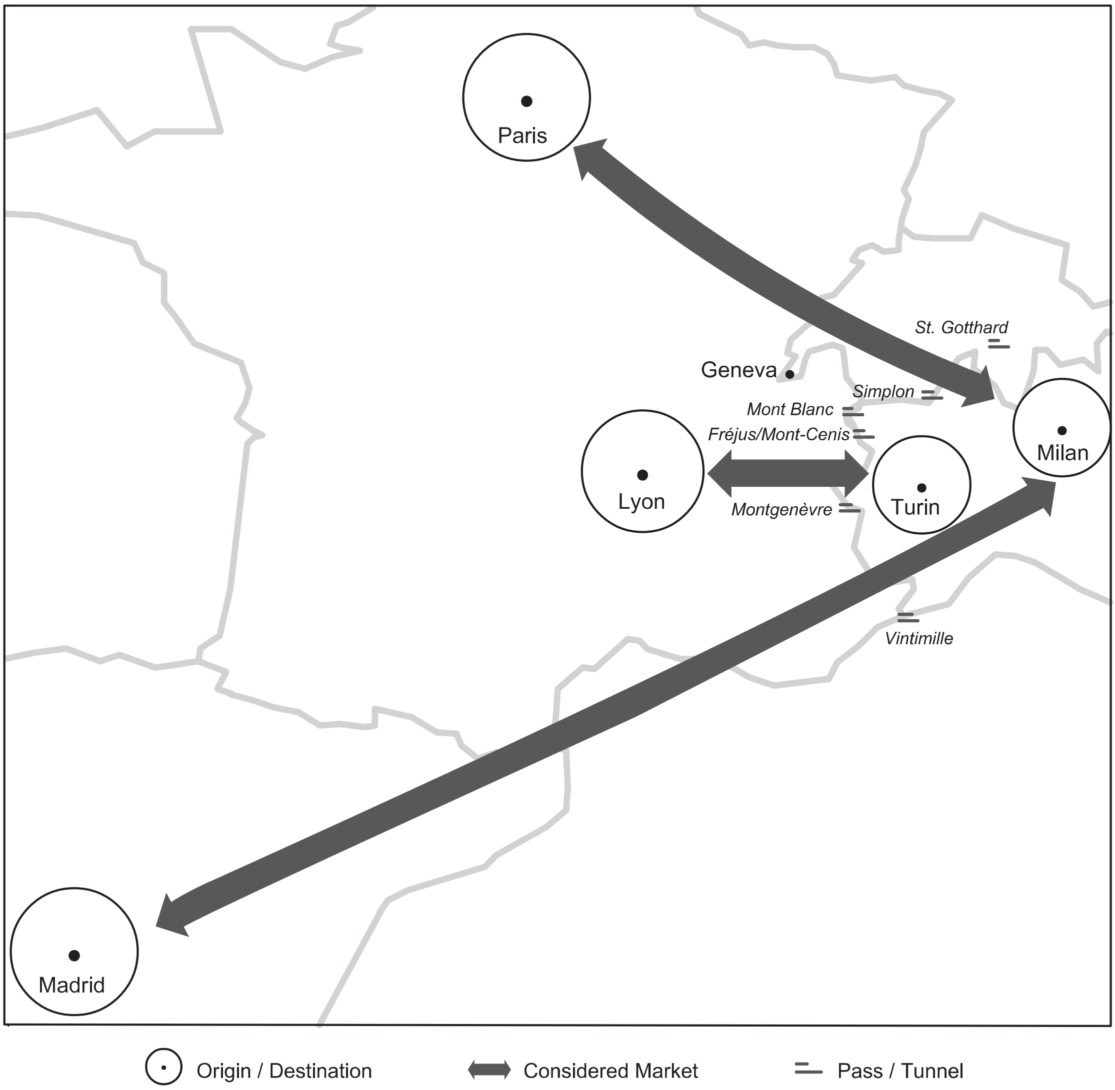

Figure 1: Schematic illustration of three considered markets

with σ being the degree of correlation between alternatives j belonging to the same group g. A high σ implies shippers give a higher weight to the group than to the alternative itself when they pick one. Competition is then fiercer between modes than between alpine paths. To be consistent with the random utility maximization concept, the parameter σ must lie between 0 and 1. In the extreme case of symmetric competition where the assumption of Independence of Irrelevant Alternatives (IIA)[11] holds between all alternatives, σ equals 0 and the model reduces to the simple logit specification. In the other extreme of segmentation, where preferences for alternatives are perfectly correlated within nests but independent between nests, σ is equal to 1.

Random components ν_{ig}, ν_{ij}, and consequently ϵ_{ij} are standard extreme value distributed.
We assume the mean utility level to be:

$$V_j = \Psi_j - h\, p_j \tag{3}$$

where:

- Ψ_j is the aggregate measure of quality of product j
- h represents the sensitivity of utility to price, that is the marginal utility of cost saving for the shipper.

We then compute the aggregate measure of quality as the weighted sum of the alternatives' characteristics:

$$\Psi_j = \alpha_1\, punctuality_j + \alpha_2\, alt_j + \alpha_3\, traveltime_j + \alpha_4\, maxcapa_j \tag{4}$$

where:

- $punctuality_j$ = ratio of highway kilometers over total road kilometers and actual published punctuality figures[12] for road and rail, respectively; expected positive sign for α_1.
- alt_j = altitude, in meters, of path j; expected negative sign for α_2.
- $traveltime_j$; expected negative sign for α_3.
- $maxcapa_j$ = maximum capacity per unit of transportation; expected positive sign for α_4.

[11] See McFadden (1981).

[12] See SBB Cargo's Annual Report (2004, page 18) and at SNCF: http://fret.sncf.com/fr/quisnous/actu/2007/presse/do070618.pdf

We attach specific values to these variables, respective to each product. Some values deserve further explanation.
Variable $traveltime_j$ includes compulsory drivers' breaks. Every 4 hours and half, a truck driver has to rest for 45 minutes; after 9 hours of driving, a truck driver has to stop and rest for 10 hours.
Variable $punctuality_j$ codes for the probability for a freight carrier to meet his travel time target. This is a reliability measure. Even though obviously correlated, variables $punctuality_j$ and $traveltime_j$ do not exactly capture the same path features.
Variable $maxcapa_j$ indicates the maximum tonnage one unit load can carry. This takes into account the fact that there is a more strict heavy goods weight constraint on using road transport as compared to rail. Finally, variable alt_j mainly codes for changing conditions of mountainous weather.

We assume these four speed and reliability measures to be the most relevant for shippers in order to assess quality of available products. Shipper i chooses the utility-maximizing alternative j, satisfying:

$$U_{ij} \geq U_{ik} \quad \forall \quad k \neq j \tag{5}$$

Normalizing the mean utility of the outside good to zero, we compute the probability of choosing alternative j from the probability of choosing group g and the probability of choosing alternative j conditional on choosing group g. We apply the methodology proposed by Berry (1994) and widely used in the estimation of differentiated products demand.[13] This methodology builds upon the assumption that observed aggregate market shares are valid approximations of choice probabilities. It allows us to derive the mean utility levels as follows:

$$\ln s_j - \ln s_0 = \Psi_j - h\, p_j + \sigma\, \ln s_{j/g} \tag{6}$$

with s_j and $s_{j/g}$ respectively being the total market share and the group market share of alternative j.
Finally, the own price elasticity of demand of the alternative j is:

$$\mu_j = h\, p_j \left[s_j - \frac{1}{1-\sigma} + \frac{\sigma}{1-\sigma}\, s_{j/g} \right] \quad \forall \quad j \in g \tag{7}$$

2.2 Supply side

We focus on the competitive aspect of cross-alpine freight transport. Competing freight carriers offer shippers a differentiated product combining a transport mode with a specific alpine tunnel or pass – Mont Blanc, Frejus, Montgenevre, Vintimille, Mont-Cenis or Gotthard (see Figure 2.2).[14]

[13] See, for example, Ackerberg et al. (2007) and Ortúzar (2001).
[14] We consider the Gotthard passage only on the transit market between Ile-de-France-Nord and Lombardia. Its road and rail market shares are sufficiently high (respectively 5.9 % and 2.0 % of

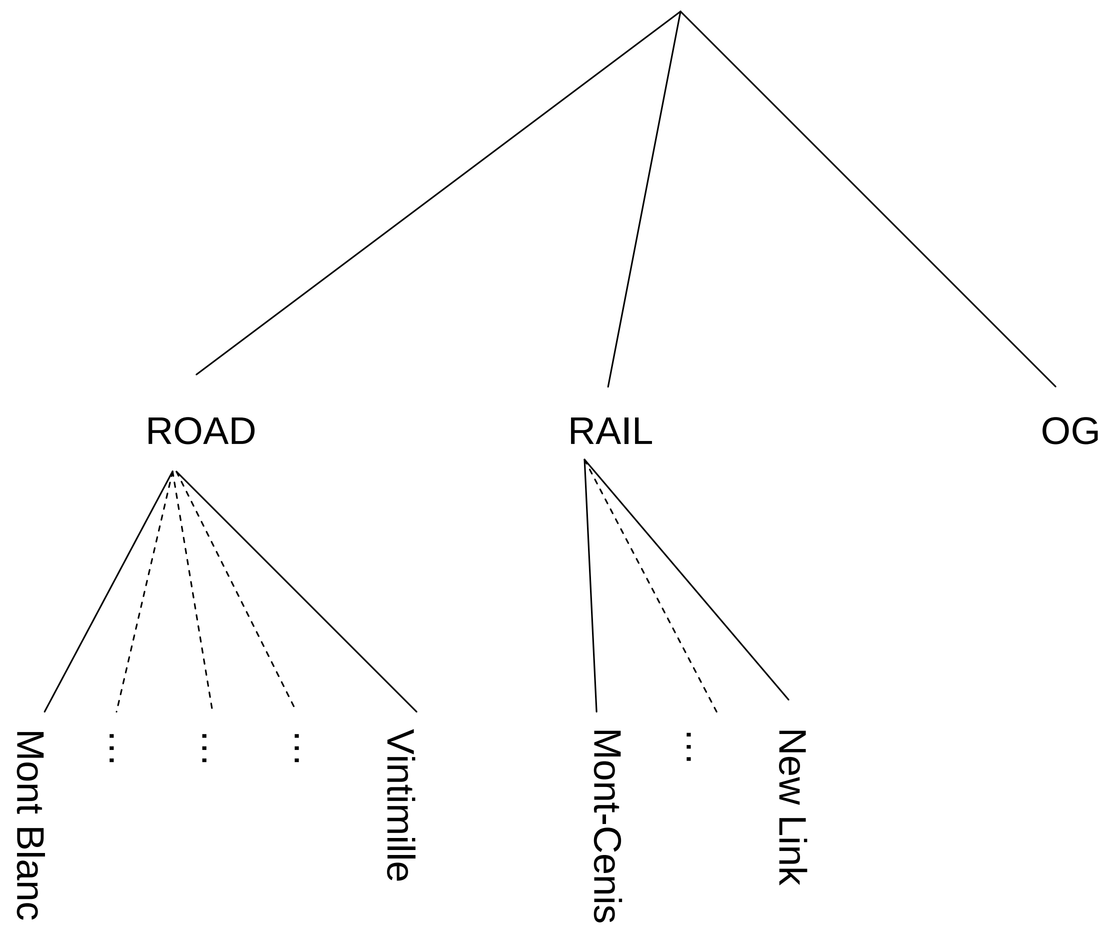

Figure 2: Basic discrete choice model

In 2004, the 'Autoroute Ferroviaire Alpine', a joint venture between SNCF and TRENITALIA providing a rail shuttle service for lorries and semi-trailers through the Fréjus tunnel, has been experimented and its related traffic reported in our data. However, the generated traffic was so low and the restrictions so numerous that we do not consider this alternative relevant for our analysis. It remains to be seen if this mode of transport will prove successful in the long run.

We assume that each differentiated product is offered by one firm only. This simplifies reality to a large extent. In particular, the road freight industry is quite atomistic. In the aggregate, 77.6% of road freight carriers employed 0 to 5 people and 2.3% of all transport companies had more than 50 employees in France in 2000. In terms of revenue, freight carriers with less than 50 employees accounted for 59.4%

their group market) for these two products to be considered as relevant competitive alternatives on this market.

of the industry's total revenue.[15] Given this structure of road supply, competition is likely to be fierce among road carriers. Therefore, we would intuitively think the latter to be price-takers rather than makers. Notwithstanding, we argue that demand is rather captive on each geographical market. This justifies to some extent the strategic role of road carriers as price-makers. In what follows, we assume that road carriers have some power to set prices above marginal cost.[16]

In equilibrium, cross-alpine freight carriers set transport prices in order to maximize their profits, knowing their competitors do the same:

$$Max \quad \Pi_j = (p_j - - c_j)\ q_j - - K \tag{8}$$

with fixed costs K.

The outcome is defined by the set of J necessary first order conditions, from Ivaldi and Verboven (2005):

$$p_j = c_j + \frac{1-\sigma}{h\left(1 - \sigma\, s_{j/g} - (1-\sigma)\, s_j\right)} \tag{9}$$

The price of a product j is therefore the sum of its marginal cost, c_j, and a mark-up term.

3 Empirical Analysis and Results

Shippers' and freight carriers' actions depend on their sensitivities to changes in the alternatives' characteristics. We want to measure the impact the prospective rail link between Lyon and Turin will have on the equilibrium market shares, prices and consumer surplus. Before rushing into the simulation analysis, we need to derive the equilibrium features of our three markets, where shippers can choose only from current available alternatives. Our data, reported in Appendix A, prevents straight statistical estimation of the model. Therefore, we need to calibrate the model, that is to find the equilibrium values of the demand parameters, h and σ, as well as the quality parameters in Equation 4. We can then define the equilibrium outcome and simulate likely changes following the new link's introduction. All numerical computations were done using *Matlab*.

3.1 Model Calibration

Following the procedure by Ivaldi and Vibes (2008), we first derive the demand parameters h and σ. To do so, we linearize Equation 7 defining price elasticities.

[15]EUROSTAF (2003), page 4.

[16]Ivaldi (2007) maintains this assumption analogously.

Table 1: Equilibrium outcomes

		Short-distance			Transit North-South			Transit East-West		
Share of outside alternative in%		15	30	45	15	30	45	15	30	45
Road Market Shares in %	Mont-Blanc	7.0	5.8	4.5	34.5	28.5	22.4	-	-	-
	Frejus	57.4	47.2	37.2	34.6	28.5	22.4	0.8	0.6	0.5
	Montgenevre	1.3	1.1	0.9	0.4	0.3	0.3	0.04	0.03	0.02
	Vintimille	0.5	0.4	0.3	0.8	0.6	0.5	83.76	68.97	54.18
	Gotthard	-	-	-	5.0	4.1	3.2	-	-	-
Rail Market Shares in %	Mont-Cenis	18.8	15.5	12.2	8.0	6.6	5.2	0.2	0.2	0.1
	Vintimille	-	-	-	-	-	-	0.2	0.2	0.2
	Gotthard	-	-	-	1.7	1.4	1.1	-	-	-
Marginal utility of cost saving (Parameter h)		0.006	0.005	0.005	0.001	0.001	0.001	0.002	0.002	0.001
Degree of within-group correlation (Parameter σ)		0.40	0.44	0.52	0.62	0.62	0.63	0.56	0.56	0.62
Own-Price Elasticities (Road)	Mont-Blanc	-4.27	-4.11	-4.19	-1.81	-1.85	-1.89	-	-	-
	Frejus	-1.48	-1.63	-1.73	-1.87	-1.92	-1.96	-6.93	-6.93	-6.78
	Montgenevre	-3.70	-3.54	-3.61	-3.25	-3.21	-3.2	-6.93	-6.93	-6.78
	Vintimille	-5.86	-5.60	-5.71	-4.34	-4.28	-4.22	-0.53	-0.98	-1.20
	Gotthard	-	-	-	-2.99	-2.96	-2.93	-	-	-
Own-Price Elasticities (Rail)	Mont-Cenis	-1.63	-1.53	-1.38	-1.23	-1.22	-1.21	-3.57	-3.58	-3.36
	Vintimille	-	-	-	-	-	-	-3.47	-3.48	-3.25
	Gotthard	-	-	-	-2.37	-2.34	-2.31	-	-	-
Consumer Surplus		296	160	43.8	1556	780	190	1092	528	150

Notes: Market shares are computed from aggregate CAFT 2004 data. All other parameters result from our calibrated model.

We do not have data on elasticities but we do have data on market shares, prices, and, contrary to Ivaldi and Vibes (2008), marginal costs. We repeatedly draw 1000 vectors of elasticities, based on the normal distribution function, with a standard deviation of 4. We set the mean of this distribution to be equal to the commodity-specific values presented in Oum et al. (1990), weighted by the commodity shares transported on each link. We then obtain values of h and σ to each draw of 1000 elasticity vectors using ordinary least squares estimation. These values allow us to derive a marginal cost vector from Equation 9 for each of these draws. We finally keep the h and σ values, as well as the associated vector of elasticities μ_j, which correspond to the closest match of predicted marginal costs with our vector of observed marginal costs. We use these results to solve the system of equations defining first the quality parameters Ψ_j, second the quality components described in Section 2.1 and Appendix A.1, and third their coefficients in Equation 4. We thus solve a system of five linear equations and four unkowns, and derive consumers' valuations of each quality variable. We present the results on elasticity and parameter values in Table 1. Note that the elasticity values are calibrated to our very specific markets and thus cannot be easily interpreted outside of these markets nor be directly compared to estimates in the literature that stem from different markets or are aggregate averages.

Small values of h underline the intuitive fact that individuals have a larger marginal utility of income than a firm's marginal utility of saving costs. On both transit markets, h is even lower. Indeed, as pointed out in section 2, long-distance freight shippers are rather large companies while short- and medium-distance freight shippers tend to be relatively small companies. Intuitively, it makes sense that the latter care more about cost savings.

The high value of σ shows low substitutability between the nests of differentiated products in the alpine freight transport market. The **mode choice** remains the main component of the shippers' decision.

In terms of market shares, only the transit market between Spain and Lombardia exhibits a true dominant alternative: Vintimille (road). This remains true whatever the OG market share. This demand rigidity is also reflected in the price-elasticities ranking on this specific market since the lowest price-elasticity is associated with the Vintimille (road) alternative. More generally, the model endogenously implies that small market shares go along with high absolute values of elasticities. Therefore market shares' rankings on our three markets directly translate into elasticities' rankings, in absolute values. We present values of the quality indices and the weights of the quality indices' components in Table 2. All coefficients have the expected signs.

As we do not know the exact market configuration, we allow the OG share to vary between 15 and 45% of the total freight transport market. Alpine traffic forecasts could help to determine the situation we are in. However, these forecasts differ across scenarios and do not appear reliable. Indeed, most of them are suspected to largely overestimate freight alpine traffic by the years 2020 and 2030.[17] In what follows we focus on the "$OG = 15\%$" case for our equilibrium and simulation analysis. Qualitatively, the results carry over to the two scenarios with larger OG market shares.

[17] ECMT Report (2001)

Table 2: Quality Indices and quality component weights

		Short-distance			Transit North-South			Transit East-West		
Share of outside alternative in %		15	30	45	15	30	45	15	30	45
Quality Index – Road	Mont-Blanc	2.93	1.84	1.05	2.50	1.58	0.91	-	-	-
	Frejus	4.27	3.10	2.12	2.54	1.62	0.95	2.74	1.86	1.01
	Montgenevre	1.39	0.42	-0.17	0.87	-0.03	-0.68	1.38	0.49	-0.17
	Vintimille	2.03	0.97	0.32	1.53	0.62	-0.05	4.73	3.87	2.73
	Gotthard	-	-	-	1.79	0.88	0.22	-	-	-
Quality Index – Rail	Mont-Cenis	2.24	1.15	0.27	0.51	-0.40	-1.08	-1.71	-2.59	-3.54
	Vintimille	-	-	-	-	-	-	-1.68	-2.55	-3.51
	Gotthard	-	-	-	-0.09	-1.00	-1.66	-	-	-
Quality Index – New Transalpine Rail Link		3.38	2.44	1.47	1.35	0.51	0.12	1.79	1.06	-0.41
Quality component coefficients	Altitude	-0.0004	-0.0006	-0.0007	-0.0001	-0.0001	-0.0002	0	0	0
	Travel time	-0.17	-0.18	-0.16	-0.24	-0.25	-0.26	-0.71	-0.74	-0.64
	Punctuality	5.49	4.59	3.68	8.28	7.72	7.25	31.31	31.78	27.23
	Max capacity	0.0076	0.0070	0.0008	0.0102	0.0086	0.0069	0.1423	0.1506	0.1089

Notes: All results are computed using the equilibrium parameters from our calibrated model.

3.2 Simulations and Results

We can now simulate the entry of the new rail link using its quality characteristics and the demand parameters from the calibrated model. Appendix C elaborates this procedure in more detail.

3.2.1 Results in the short-distance market between Lyon and Turin

Table 3 shows results for different initial market shares of the outside alternative in point-to-point transport between Lyon and Turin.[18]

A global overview of our first simulation result underlines three important effects of the new link provision in the regional market. First, the two rail alternatives manage to capture more than 37% of the total inter-regional traffic, the new transalpine link taking most of it (27%). Second, the two incumbent Fréjus products (road and rail) lose the most after the introduction of the new link. Third, consumer surplus increases by almost 6%.

Table 3: Introduction of the new transalpine rail link, short-distance

		Short-distance (Lyon-Turin)					
Initial share of outside alternative in %		15		30		45	
Values (in euro) and Change in %		Value	Δ	Value	Δ	Value	Δ
Road prices	Mont-Blanc	481	+12.6%	486	+2.3%	483	+1.7%
	Frejus	555	+13.5%	561	+14.7%	556	+13.7%
	Montgenevre	239	-37.6%	243	-36.5%	241	-37.1%
	Vintimille	398	-33.6%	403	-33.0%	401	-33.2%
Rail prices	Mont-Cenis	343	0.0%	344	+0.3%	339	-1.2%
	New Transalpine Link	434	-	460	-	462	-
Road market shares in %	Mont-Blanc	7.1	-4.0%	5.8	0.0%	4.8	+6.6%
	Frejus	32.0	-44.6%	26.8	-43.2%	22.4	-40.0%
	Montgenevre	5.8	+346.0%	4.5	+310.0%	3.8	+322.0%
	Vintimille	3.5	+600.0%	2.7	+575.0%	2.3	+666.0%
Rail market shares in %	Mont-Cenis	10.2	-45.7%	7.5	-51.6%	5.1	-58.2%
	New Transalpine Link	27.6	-	25.1	-	19.6	-
Market share of outside alternative in %		13.8	- 8.0%	27.7	-7.6%	42.1	-6.4%
Consumer surplus		313	+ 6.0%	181	+13.0%	69.6	+60.0%

The rail mode plays a very important part and appears fairly competitive on the inter-regional market between Lyon and Turin. Facing the biggest loss in terms of market shares, the representative supplier of the Frejus road product nevertheless increases its prices by 13.5%. Its historical competitors tend to align their prices around an average of 365 euros.

Going into detail, we see that prices do not vary homogenously. The two main road alternatives, Fréjus and Mont-Blanc, increase their prices by similar amounts:

[18] The new alternative is assumed to have the same marginal cost as the historical *Mont-Cenis* alternative served by SNCF since we do not have expected cost data for the new alternative. In principle, implementing a variation in cost would be straightforward.

13.5% and 12.6% respectively. By contrast, the two "outsider" road products, Montgenèvre and Vintimille, cut their prices by quite large amounts: 37.6% and 33.6% respectively.

The rather captive Frejus demand[19] and its high quality index relative to the other alternatives explain most of the price reaction to the new rail product. Indeed, these two features allow the Frejus road supplier to compensate its loss in market share - mostly to the benefit of the new rail alternative - by a price increase. The less competitive Montgenèvre road and Vintimille road providers benefit from the weaker market position of the Frejus road alternative. They even compete more fiercely[20] to gain market shares at the expense of the Frejus road supplier. Therefore, road carriers' reactions to the entry of the new "Liaison Ferroviaire Lyon-Turin" depend on their relative historic market power. Historic dominant providers are very sensitive to the induced inter-modal competition and consecutively alter their pricing behavior to compensate their loss in market shares. Historic "outsiders" take advantage of their competitors' weaker posture and toughen their pricing strategies in order to get "a bigger piece of the bigger cake".

The substantial increase in the rail market share is partly due to the increase of global market size.[21] However, most of it results from a modal shift in favor of the new rail alternative. Undoubtedly, the strong market position of the new link comes at the expense of the historical rail alternative Mont-Cenis. The latter, however, manages to keep a reasonable share of total regional traffic, thanks to its price competitiveness relative to the two main road products. As for the new link, its high quality index as well as its competitive price make it a viable competitor of the two main road alternatives on the regional market. Note, however, that we are looking at a very specific short-distance market. We thus have to be cautious when comparing the change in market shares to ones in more geographically aggregate settings.

As a concluding remark, the introduction of a high quality alternative and the decrease in the OG's market share induce an improvement of consumer surplus. However, this consumer gain remains of rather low magnitude.

3.2.2 Results in the transit market between Ile-de-France and Lombardia

Table 4 shows the results for freight journeys between the Ile-de-France and Lombardia regions.

Three salient facts summarize behaviors on this North-South freight transit market. First, intra-modal competition within both nests is fierce. Second, the new rail

[19] See the own price elasticity $\mu_{Frejus} = -1.48$.

[20] There are no tunnel fees for these passages. Hence, substantially lower marginal costs allow significant price drops.

[21] See the OG share decrease of 8%.

alternative captures 11.2% of the total market, quasi as much as the Gotthard (11.4%) which ranks third among road products. Third, consumer surplus increases by 6%.

Table 4: Introduction of the new transalpine rail link, North-South

		Transit North-South (Ile-de-France-Lombardia)					
Initial share of outside alternative in %		15		30		45	
Values (in euro) and Change in %		Value	Δ	Value	Δ	Value	Δ
Road prices	Mont-Blanc	1097	+4.0%	1101	+4.2%	1095	+3.7%
	Frejus	1121	+2.5%	1125	+3.0%	1118	+2.3%
	Montgenevre	704	-37.0%	709	-36.3%	713	-36.0%
	Vintimille	843	-43.5%	847	-43.3%	849	-43.1%
	Gotthard	754	-30.5%	757	-30.2%	757	-30.2%
Rail prices	Mont-Cenis	916	-0.4%	916	-0.4%	917	-0.3%
	Gotthard	868	-5.0%	872	-5.5%	876	-4.0%
	New Transalpine Link	1231	-	1269	-	1294	-
Road market shares in %	Mont-Blanc	26.6	-23.0%	22.2	-22.1%	17.9	-20.1%
	Frejus	27.7	-20.0%	23.1	-19.0%	18.6	-17.0%
	Montgenevre	1.2	+200.0%	1.0	+233.0%	0.7	+133.0%
	Vintimille	4.4	+450.0%	3.6	+500.0%	2.7	+440.0%
	Gotthard	11.4	+128.0%	9.4	+129.0%	7.3	+128.0%
Rail market shares in %	Mont-Cenis	3.1	-61.2%	2.4	-63.6%	1.8	-65.4%
	Gotthard	0.7	-59.0%	0.6	-57.1%	0.4	-63.6%
	New Transalpine Link	11.2	-	9.9	-	8.2	-
Market share of outside alternative in %		13.7	-8.0%	27.9	-7.0%	42.3	-6.0%
Consumer surplus		1649	+6.0%	876	+12.3%	295	+55.3%

As in the regional market, the entry of a higher quality rail product induces "predatory" behavior among outsiders, namely Montgenèvre, Vintimille and Gotthard road path providers, within the road nest. The latter engage in large price-cuts – -37%, -43.5%, -30.5% respectively – in order to increase their market share – by 200%, 450% and 128% respectively – and take the best out of weakened dominant Fréjus and Mont-Blanc road suppliers. However, contrast with the regional market, the new alternative does not trigger noticeable modal shift. It succeeds in attracting new shippers in the market – 8% decrease in the OG share – but globally fails in capturing demand from road alternatives. Its higher quality makes it the best rail alternative despite a price even higher - 1231 euros against 1121 and 1097 euros - than the ones of its two major road competitors, Mont-Blanc and Fréjus. Nevertheless, the new link does not appear to be competitive relative to road supply. In this respect the "Liaison Ferroviaire Lyon-Turin" alone cannot be the relevant modal shift device its proponents claim it to be. At least not on the North-South transit freight market. Therefore, only a more global transport policy scheme taking into account the strategic behavior in both supply and demand may achieve a substantial shift in transport modes.

Albeit of absolute importance, induced variations in prices and market shares do not alter the historic relative ranking of the different alpine products.

3.2.3 Results in the transit market between Spain and Lombardia

In the transit market between Spain and Italy shippers use the rail path in Vintimille. In Table 5 we see that the new rail link does not significantly impact global demand for shipping between Spain and Lombardia. Indeed, the decrease in the OG market share only amounts to 0.5% and consumer surplus does not change at all. Therefore, we conjecture that the new alternative's 7% market share corresponds to a genuine modal shift on this transit market. Indeed, the new link erodes road alternatives' market edge of 6,5%. Add to this the 0.5% decrease in the outside good share, it almost amounts to the total rail alternatives' share.

Table 5: Introduction of the new transalpine rail link, East-West

		Transit East-West (Spain-Lombardia)					
Initial share of outside alternative in %		15		30		45	
Values (in euro) and Change in %		Value	Δ	Value	Δ	Value	Δ
Road prices	Frejus	1165	-39.5%	1159	-40.0%	1158	-40.0%
	Montgenevre	893	-53.2%	893	-53.2%	898	-53.0%
	Vintimille	1995	+5.5%	1962	+3.7%	1941	+2.6%
Rail prices	Mont-Cenis	1189	-12.0%	1189	-12.0%	1195	-12.7%
	Vintimille	1173	-13.2%	1173	-13.2%	1179	-12.7%
	New Transalpine Link	1601	-	1598	-	1691	-
Road market shares in %	Frejus	13.6	+1600.0%	10.4	+1633.0%	7.6	+1420.0%
	Montgenevre	1.6	+3900.0%	1.2	+3900.0%	0.85	+4150.0%
	Vintimille	62.88	-25.0%	53.2	-22.8%	44.65	-17.6%
Rail market shares in %	Mont-Cenis	0.01	-95.0%	0.007	-96.5%	0.005	-95.0%
	Vintimille	0.01	-95.0%	0.008	-96.0%	0.005	-97.5%
	New Transalpine Link	7.0	-	6.4	-	3.02	-
Market share of outside alternative in %		14.92	-0.5%	28.785	-0.4%	43.87	-2.5%
Consumer surplus		1092	0.0%	565	+7.0%	184	+22.6%

Far more striking, however, is the gain of the Frejus road alternative which succeeds in elevating its market share by 12.8 percentage points due to strong price competition. In this market, the latter alternative takes the outsider role benefitting most from competitive price-setting on the dominant Vintimille alternative given its extremely captive demand.

3.2.4 Results from a change in the cost structure in the transit market between Ile-de-France and Lombardia

Table 6 shows results of a change in the cost structure in freight journeys between the Ile-de-France region and Lombardia. We simulate a twofold increase of tunnel fees at the Mont-Blanc and Frejus road tunnels and a simultaneous reduction of marginal costs of rail transport by one half. We interpret this as a political measure of cross-subsidizing from road to rail transport. The change in the cost structure can achieve a modal shift comparable to that induced by the introduction of an entirely new infrastructure.

Comparing Tables 4 and 6 reveals a comparable reduction of the main French passages' market shares by raising their tunnel fees. An important part of this reduction is absorbed by the Gotthard road alternative tripling its market share due to its cost advantage. Modal shift from road to rail is not as strong - 11.1% rail share

Table 6: Increase in road tunnel fees and reduction of rail marginal costs

		Transit North-South (Ile-de-France-Lombardia)					
Initial share of outside alternative in %		15		30		45	
Values (in euro) and Change in %		Value	Δ	Value	Δ	Value	Δ
Road prices	Mont-Blanc	1275	+20.7%	1278	+21.1%	1272	+20.5%
	Frejus	1297	+18.7%	1300	+19.0%	1294	+18.4%
	Montgenevre	707	-36.6%	711	-36.2%	715	-35.8%
	Vintimille	853	-42.9%	857	-42.6%	858	-42.5%
	Gotthard	783	-27.8%	785	-27.6%	782	-27.9%
Rail prices	Mont-Cenis	916	-0.4%	925	+0.6%	931	+1.2%
	Gotthard	682	-25.3%	688	-24.7%	693	-24.1%
Road market shares in %	Mont-Blanc	23.9	-31.1%	19.5	-31.6%	15.3	-31.5%
	Frejus	24.9	-28.1%	20.4	-28.6%	16.0	-28.4%
	Montgenevre	1.8	+331.8%	1.4	+317.1%	1.0	+296.2%
	Vintimille	6.4	+753.6%	5.1	+719.3%	3.8	+676.7%
	Gotthard	15.7	217.1%	12.6	+209.5%	9.7	+201.3%
Rail market shares in %	Mont-Cenis	7.9	-1.0%	6.3	-3.3%	4.9	-5.5%
	Gotthard	3.2	+92.6%	2.6	+88.2%	2.0	+82.8%
Market share of outside alternative in %		16.2	+8.0%	32.1	+7.0%	47.3	+5.1%
Consumer surplus		1471	-5.5%	690	-11.5%	103	-45.8%

as opposed to 15.0% when introducing the new alternative - and only present towards the Gotthard rail passage. While the introduction of a new rail alternative induces new traffic, the fact of higher road costs and lower rail costs increases the share of the outside good. Given the absence of a new high-quality alternative, this political measure therefore reduces total traffic as well as consumer surplus. The relevant question is then whether the difference in consumer gains/losses over-compensates the costs of building a new infrastructure. At the very least, this simulation exercise illustrates that creating new infrastructure may not be an exclusive solution but that there exist various alternative policy measures leading to a comparable end.

4 Conclusion

The model used in this paper allows to derive demand and supply equilibrium behavior in a market with product differentiation. We apply this model to the alpine freight transport market with differentiated "mode & alpine path" products in order to test the competitive viability of the prospective "Liaison Ferroviaire Lyon-Turin" project.

As a first structural result we find limited substitutability between freight transport products on a North-South transit axis, despite their heterogeneity beyond the mere modal split. Indeed, on our Ile-de-France-Lombardia market, the shippers' decision remains largely based on mode choice. The new high-quality rail alternative attracts new demand but does not succeed in reducing road demand. When we shorten this axis to the regional market between Lyon and Turin, both a modal shift and an increase in demand for shipping occur, securing *exactly the same* variations in OG market share and consumer surplus as in the North-South market. Therefore, the global impact of the new link on the latter transit market seems to be driven by local factors. The demand rigidity for road raises a methodological problem: mean utility specification demands a profound knowledge of shippers' choice criteria.[22] Micro-level data – collected during face-to-face interviews for instance – would be of great help in this respect. Precise criteria relevant for modal shift could be revealed this way and appropriate policy measures undertaken.

In contrast with the North-South analysis, the West-East transit market appears a better candidate for modal-shift. Between Spain and Lombardia indeed, the new rail link appears attractive enough for shippers to switch modes. Note that global traffic does not increase after the introduction of the new link, suggesting higher volatility of shippers' preferences on this transit axis. Should European rail integration be fostered, the new transalpine link between Lyon and Turin could play a complementary part among other urgent projects. In this respect, it would be of interest to compare respective impacts on the West-East transit axis of the Lyon-Turin Transalpine and the Perpignan-Figueras Transpyrenees.[23]

From a modeling viewpoint, improving the approximation of product flexibility – so far captured by variable *"Punctuality"* in our mean utility specification – should receive particular attention in future studies on the subject. As a matter of fact we believe the most obvious drawback of rail freight transport is its *exclusivity*: choosing rail in Lyon excludes changing modes until Turin. A delay forecast after the train departure cannot yet find remedy in a switch to a more flexible transport mode. In this respect inter-modality seems to be the key component of a competitive rail product. So far, however, inter-modal freight transport has not had the success needed for a significant modal shift. In the French Alpes, the "Rolling Highway" has been inexistent until 2005 when it accounted for 0.7% of total French Alpine freight tonnage. The corresponding value for Switzerland in 2005 was 5.2%. While on the rise in Switzerland, unaccompanied combined freight tonnage has been falling in the French Alpes from 8.6% in 2000 to 5.9% in 2005, and this even while observing a decrease in total freight tonnage.[24]

Based on our analysis, the construction of a new high quality infrastructure may only be one tool out of a global modal shift-oriented policy toolbox. For the French Alpine corridor, more direct and committed intervention based on a variety of policy measures as observed in Switzerland may open a more fruitful path to the political goal of increasing modal shift towards rail.

[22] Time and monetary costs certainly remain the most important.

[23] http://www.nouvelletraverseedespyrenees.com/historique.html

[24] Bundesamt für Verkehr (2006)

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Appendix A: Data

A.1 Supply side

In order to characterize the alternatives and accurately assess freight carriers' costs and posted prices, we conducted several phone interviews and gathered all available (to us) information.

However, we cannot directly observe marginal cost. We collected data on costs and prices of infrastructure use, and on fuel consumption to approximate them. These components are short-run cost variables. That is why we deliberately leave aside personnel costs which correspond to long-run costs. This choice of components seems all the more reasonable that our cost values are very close to the ones computed and published by the Comité National Routier, on an annual basis.[25] We explain our approximations and computations in Appendix B. In what follows we detail our price and marginal cost approximations on each market.

A.1.1 Prices, costs, and characteristics for the short-distance market

Table 7: Prices and Marginal Cost, Lyon - Turin, 24t load (in Euros)

		MC	Price
Road	Mont-Blanc	365	475
Road	Frejus	331	489
Road	Montgenevre	126	383
Road	Vintimille	290	600
Rail	Mont-Cenis	216	343
Rail	New transalpine link	216	reported after simulation

[25] http://www.cnr.fr/grilles_couts/e-docs/00/00/00/26/document_grille_cout.phtml

Table 7 shows collected prices and computed marginal cost for each passage and mode. Table 8 presents the above-described quality components of the 5 existing alternatives as well as the new transalpine link. Travel times in hours are calculated based on speed, distance and stopping periods.

Table 8: Alternative characteristics between Lyon and Turin

		Altitude	Punctuality	Travel Time	Capacity
Road	Mont-Blanc	1328	0.87	6	24
Road	Frejus	1158	0.89	4	24
Road	Montgenevre	1860	0.52	6	24
Road	Vintimille	9	0.97	21	24
Rail	Mont-Cenis	1158	0.77	11.60	60
Rail	New transalpine link	478	0.80	7.60	60

A.1.2 Prices, costs, and characteristics for the transit market between Ile-de-France-Nord and Lombardia

Table 9 shows prices and marginal cost for each product on the Ile-de-France-Nord-Lombardia transit market. Given their significant market shares, we include two Swiss alternatives on this North-South freight market: Gotthard-rail and Gotthard-road.

Table 9: Prices and Marginal Cost, Ile-de-France-Nord - Lombardia, 24t load (in Euros)

		MC	Price
Road	Gotthard	348	1085
Road	Mont-Blanc	553.5	1056
Road	Frejus	564.5	1093
Road	Montgenevre	357	1114
Road	Vintimille	479	1493
Rail	Mont-Cenis	518	920
Rail	Gotthard	515	913
Rail	New transalpine link	518	reported after simulation

In Table 10, we observe that the Swiss alternatives are competitive in terms of travel time and punctuality. This is as expected, of course, as we are looking at North-South transit. Intuitively, we would suggest that location matters most for the attractiveness of transport infrastructure.

Table 10: Alternative characteristics between Ile-de-France-Nord and Lombardia

		Altitude	Punctuality	Travel Time	Capacity
Road	Gotthard	1150	0.98	23	24
Road	Mont-Blanc	1328	0.94	23	24
Road	Frejus	1158	0.96	23	24
Road	Montgenevre	1860	0.83	25	24
Road	Vintimille	9	0.98	28	24
Rail	Mont-Cenis	1158	0.77	28	60
Rail	Gotthard	1150	0.79	28	60
Rail	New transalpine link	478	0.80	24	60

A.1.3 Prices, costs, and characteristics for the transit market between Spain and Lombardia

For values on the transit market we use "geographic averages", that is we compute price and cost values for the O-D relationship Madrid — Lombardia (North of Italy) as a proxy. This is a simplification we need to make given our modeling framework but we argue that these averages will be informative nonetheless. A more complex analysis with far better data would be required to model intra-European transit in significantly more detail. Table 11 shows prices and marginal cost for each passage and mode in Spain-Italy transit.

Table 11: Prices and Marginal Cost, Spain - Lombardia, 24t load (in Euros)

		MC	Price
Road	Frejus	832	1926
Road	Montgen6̀12	586	1910
Road	Vintimille	606	1891
Rail	Mont-Cenis	913	1351
Rail	Vintimille	897	1351
Rail	New transalpine link	913	reported after simulation

Table 12 presents the quality components of the 6 existing alternatives as well as the new transalpine link.

A.2 Demand side

Within our random utility framework and to best assess the benefits from the planned Lyon-Turin new link, individual level data would be needed. This kind of micro-level data can be produced through expensive and time-consuming surveys that are not feasible in the scope of this paper. We therefore choose to use the inversion method proposed by Berry (1994). This methodology requires aggregated data, such as market shares and information on prices, along with some quality variables within the discrete choice

Table 12: Alternative characteristics between Spain and Lombardia

		Altitude	Punctuality	Travel Time	Capacity
Road	Mont-Blanc	1328	0.97	44	24
Road	Frejus	1158	0.97	43	24
Road	Montgenevre	1860	0.91	43	24
Road	Vintimille	9	0.99	42.5	24
Rail	Mont-Cenis	1158	0.77	49	60
Rail	Vintimille	9	0.77	48	60
Rail	New transalpine link	478	0.80	45	60

framework. Applying this method on the O-D pair Lyon-Turin we follow Ivaldi and Vibes (2008) and look for market shares of passages on this particular link.

For all three markets, we obtain market shares based on tons transported on each alpine passage from the Cross-Alpine Freight Transport (CAFT) survey 2004. Austrian, French, and Swiss authorities interview a representative sample of all Alp-crossing traffic every 5 years. In the CAFT survey, information on the origin, destination, alpine passage, transport mode, weight, etc. is collected. Table 13 illustrates market shares for traffic between Lyon and Turin which accounts for 4.6% of total freight traffic crossing the French-Italian Alpine corridor in 2004.

Table 13: Freight shares between Lyon and Turin, 2004

Passage		Market Share
Road	Mont-Blanc	8.2 %
Road	Frejus	67,5 %
Road	Montgenevre	1.6 %
Road	Vintimille	0.5 %
Rail	Mont-Cenis	22.2 %

Table 14 illustrates these shares for the freight transit traffic between regions Ile-de-France-Nord and Lombardia. This traffic amounts to 1.6% of total freight traffic crossing the Western alpine corridor in 2004.

Table 15 presents the passages' shares for transit traffic between Spain and Lombardia. This market accounts for 6.4% of all traffic crossing the French-Italian Alpine corridor in 2004. On this O-D relation, the Vintimille-road product clearly dominates the market. Strikingly enough are the comparable shares of both rail products: Mont-Cenis-rail captures a market share almost as large as the one of its Vintimille-rail group-competing product. Therefore, a new and better performing rail link close to the geographical location of the Mont-Cenis tunnel may be able to capture some market share.

The markets in our analysis cover 12.6% of total freight traffic reported in the CAFT 2004 database. One may be tempted to question the relevance of these markets or their

Table 14: Freight shares between Ile-de-France-Nord and Lombardia, 2004

Passage		Market Share
Road	Gotthard	5.9%
Road	Mont-Blanc	40.7%
Road	Frejus	40.8%
Road	Montgenevre	0.4%
Road	Vintimille	0.9%
Rail	Mont-Cenis	9.3%
Rail	Gotthard	2.0%

Table 15: Freight shares between Spain and Lombardia, 2004

Passage		Market Share
Road	Frejus	0.90 %
Road	Montgenevre	0.04 %
Road	Vintimille	98.56 %
Rail	Mont-Cenis	0.24 %
Rail	Vintimille	0.26 %

ability to capture representative behavior by consumers. As pointed out before, geographic features of one “product” are crucial to its competitiveness. The new rail link explicitly targets North-South freight traffic and aims at diverting it from other Alpine paths. The *ex ante* traffic size, on this peculiar market, is not relevant for the new rail to prove attractive or not. Moreover, we do account for market size, and its likely extension, via the outside option.

Appendix B: Cost and Price Approximations

B.3 Marginal costs - Road

For trucks, marginal costs include costs of infrastructure use, such as road and tunnel fees, and fuel costs. The former are available from infrastructure operators, that is highway and tunnel operators. We take fuel consumption values given by an online route planner.[26] Using the per liter price for truck diesel in June 2004 of 0.87 cents, we compute fuel costs

[26] http://www.autoroutes.fr

on each passage.

B.4 Prices - Road

Pricing in truck freight is mainly done according to the type of carried goods, weight and distance. These components obviously leave room for price discrimination that we cannot take into account in this study. We use prices generated by a pricing tool used by a typical road freight carrier and obtained via telephone interviews with road freight companies. For more precise results, a more sophisticated - but less tractable - price behavior, for example non-linear pricing, should be adopted.

B.5 Marginal Costs - Rail

In rail transport, marginal costs are also given by the costs of infrastructure use and fuel consumption. Data on infrastructure charges can be found either at RFF that manages and operates the French rail network, or, at the European level, at the EICIS Portal.[27] Energy consumption of a standard locomotive pulling a standard train of 800 tons[28] is considered here. We also account for the higher energy consumption on tracks that exhibit steeper slopes. As we were not able to extract values on operational costs of freight trains from several interviews with large rail freight companies, we have to use rather hypothetical values here. Again, knowing exact marginal cost values could enhance the quality of our results. Furthermore, there obviously exists a remarkable degree of heterogeneity in train technologies, train sizes and weights that we leave aside in this study for the sake of simplicity and tractability.

B.6 Prices - Rail

For rail prices, we take tariffs for a 24t shipment on a standard 4-axle train wagon with a capacity of 60t on the distance of the existing rail link from SNCF's freight tariff scheme.[29] From an interview with a representative of a large European freight carrier we know, however, that actual prices usually lie about 15% below these tariffs, due to the possibility of negotiation, quantity discounts and else.

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[27] http:// www.eicis.com
[28] Christen et al. (2004)
[29] http://fret.sncf.com/fr/espclnt/ncc/index.asp

Appendix C: Simulation of the Entry of a New Transport Link

Once the model is calibrated we can proceed to the simulation of the entry of a new alternative. Since we know the new alternative's quality characteristics and have previously derived the coefficients of quality components in the quality index we obtain the quality index for the new alternatives and therefore V_j. Next, we need to recover freight carriers' pricing behavior when a new competitor arrives. We do this using the pricing Equation 9 and the following expressions for the alternatives' market shares that incorporate the quality index in the nested logit setting (see Clerides, 2008, or Trajtenberg, 1989):

First, define:

$$D_g = \sum_{j \in J_g} \mathrm{e}^{\frac{V_j}{1-\sigma}} \tag{C.1}$$

Then, we obtain:

Intra-group market share:

$$s_{j/g} = \frac{\mathrm{e}^{\frac{V_j}{1-\sigma}}}{D_g} \tag{C.2}$$

Group market share:

$$s_g = \frac{D_g^{(1-\sigma)}}{\sum_g D_g^{(1-\sigma)}} \tag{C.3}$$

Total market share:

$$s_j = s_{j/g}\, s_g = \frac{\mathrm{e}^{\frac{V_j}{1-\sigma}}}{D_g^{\sigma}\left[\sum_g D_g^{(1-\sigma)}\right]} \tag{C.4}$$

Share of the Outside good:

$$s_0 = \frac{1}{\sum_g D_g^{(1-\sigma)}} \tag{C.5}$$

We solve Equation 9 for the new price vector p and obtain the new market shares using the above expressions, which is straight forward. Disposing of prices and mean utility values after the introduction of new alternatives we can furthermore compare the net consumer surplus the decision maker faces before and after the introduction of a new alternative. We take the expression in Ivaldi and Verboven (2001):

$$CS = \frac{1}{\alpha} ln \left(\sum_{g=1}^{G} D_g^{(1-\sigma)} \right) \tag{C.6}$$

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Chapter 2

Consumer Welfare and Unobserved Heterogeneity in Discrete Choice Models: The Value of Alpine Road Tunnels

joint with Daniel Cerquera

1 Introduction

In devising informed policy measures, reliable estimates of welfare implications for the affected individuals and firms are crucial. Measuring consumer surplus in the discrete choice framework (see McFadden, 1973), for example from changes in choice sets or in product characteristics, goes back to Small and Rosen (1981), who derive the Hicksian compensating variation in discrete choice models. Prominent examples of estimating welfare implications using this framework are Trajtenberg (1989), on the introduction of CT scanners, Goolsbee and Petrin (2002), on Satellite TV in the US, and Petrin (2004), on the introduction of the Minivan in the US automotive market. Much progress has been made in identifying and providing flexible ways of modeling unobserved consumer heterogeneity. The workhorse model is the random coefficients, or mixed, multinomial logit model introduced by Boyd and Mellman (1980) and Cardell and Dunbar (1980). The model can approximate any random utility model arbitrarily well, as shown by McFadden and Train (2000), *if* the researcher knows the correct distribution of random coefficients a priori. Just as the above examples, most applied work has imposed parametric distributions on the coefficients over which individuals are assumed to differ. Standard choices of distributions are the normal and the log-normal distribution. The main reasons for their tremendous success have been computational tractability, flexibility, and the convenient incorporation of instrumental variables by using the method of simulated moments.

In this paper, we provide insights into the impact of distributional assumptions in modeling unobserved heterogeneity on estimates of a welfare measure. If the researcher aims to estimate consumer surplus, we expect an adequate representation of the true underlying taste distribution to be important. To investigate this expectation further, we consider alternative distributional assumptions in a transalpine route choice setup. We employ revealed preference data from a large-scale transport survey in 2004 and analyze the implications of a transport policy measure which has been debated in recent years: the closure of the Mont Blanc tunnel to freight traffic. In particular, we consider the choice of road tunnels by transalpine freight traffic, where individual decision makers face a set of mutually exclusive options. We proceed by estimating three discrete choice model specifications. First, the simple fixed coeffient logit model. Second, the parametric random coefficients logit model. Third, a nonparametric estimator of preference distributions recently proposed by Bajari et al. (2007, 2010), henceforth BFKR. For each specification, we estimate the loss in consumer surplus caused by a change in choice sets and compare the results. To our knowledge, we are the first to apply the BFKR estimator to real-world data in a random coefficients logit setting. Many applications have estimated welfare effects but we are not aware of many studies exploring the role of unobserved heterogeneity in this context. A notable exception are Hynes et al. (2008), who compare welfare implications of a random coefficients logit model and a latent class model of kayakers' destination choices of whitewater sites in Ireland. They find no significant differences in welfare results from both models.

The Alps are Europe's highest and most extensive mountain range running from Mediterranean France to southern Austria. With its extreme geography, the alpine region not only forms a natural frontier between Italy and central Europe, but provides the unique gateway for ground transport between south-eastern European regions (and beyond) and central and northern Europe. Due to the limited number of crossing points, the Alps are a natural bottleneck at the core of European economic activity. After the introduction of the European single market, the opening of eastern Europe and the corresponding enlargement of commercial relations within the European Free Trade Association (EFTA), transalpine freight traffic has become not only an important topic in European politics but a key component of transport infrastructure planning. Mountainous road infrastructure exhibits elevated risks for its users, among others, due to its reliance on long underground passages. Tragic displays of this fact have been the accidents in the Mont Blanc tunnel in 1999, the Tauern tunnel in 1999, the Gotthard tunnel in 2001, and the Frejus tunnel in

2005. The most severe of these four accidents, in the Mont Blanc tunnel, cost 39 lives and lead to a full closure of the tunnel for three years. These events have brought about policy initiatives in various forms. For the alpine regions, additional investment in security measures at tunnels and formulating the objective of shifting freight to rail and maritime transport have been particular examples.[1] Proposed safety measures have reached as far as the closure of certain road tunnels to freight traffic.[2] Quantifying the short-term monetary loss incurred by the freight transport sector from the closure of a given tunnel is of interest to inform policy decisions for two main reasons. First, quantifying the monetary consequences of a potential closure for its users are at the core of any cost-benefit analysis. Second, the burden caused by unintended closures due to accidents must be known when assessing the monetary benefits of investment in safety measures. There are few studies evaluating the accidental or deliberate closure of transport infrastructure. One exception are Bilotkach et al. (2009) who exploit the collapse of a freeway interchange in the San Francisco Bay Area to analyze sensitivity of pricing behavior to demand shocks. In practice, removing transport infrastructure is not as exotic a discussion as one may think. A recent debate has been ignited by the potential closure of an Expressway in the Bronx in New York City (see Dolnick, 2010).

To investigate the impact of distributional assumptions on welfare implications, we estimate the monetary relevance of alpine road infrastructure to road freight crossing the Western Alpine corridor. In particular, we estimate the monetary loss incurred by the freight transport sector due to a hypothetical closure of the Mont Blanc tunnel. We analyze a hypothetical closure since we cannot use the actual exogenous event in 1999 for identification as point in time information is lacking in the data. Our welfare analysis should be seen in a narrow sense. Total welfare encompasses not only direct changes in consumer surplus but also changes in external effects, e.g. caused by congestion or nuisances to other travelers, as well as macroeconomic variables such as regional development or trade. Particularly in the alpine regions, estimating the total social cost would need to include both the direct costs to users (changes in consumer surplus) and external effects on non-freight users (congestion, for example) and non-users of the infrastructure such as inhabitants of the respective alpine valleys. The monetary cost of injuries, property damage and business interruption should be more directly quantifiable while macro-economic effects on economic activity and trade are difficult to identify. Here we focus on direct short-term effects likely to be central to the current political discussion.

We find that both parametric assumptions and the dimensionality of modeled unobserved heterogeneity have a significant impact on welfare results. Our BFKR estimates predict economically significantly higher annual losses in user surplus due

[1]See, for example, European Commission (2001, 2006) and Andrews (2001).

[2]See European Parliament (2001), and articles in the LA times (http://articles.latimes.com/2002/mar/10/news/mn-32111) and the BBC (http://news.bbc.co.uk/2/hi/europe/1863245.stm).

to the Mont Blanc tunnel closure. While the latter implies a loss of €5.39 Mio, the parametric random coefficients logit model predicts a loss of €2.97 Mio in specifications where both price and time are assumed to have random coefficients. With one random coefficient, the BFKR estimate is almost double that of the parametric random coefficients logit estimate, €7.09 Mio versus €3.62 Mio. Compared to both the fixed coefficients logit and the BFKR estimates, both parametric random coefficient specifications underestimate the loss in consumer surplus. We caution the exclusive use of standard distributional assumptions in modeling heterogeneity and demonstrate the simple implementation of the BFKR estimator.

The next section presents our data on freight traffic in the alpine region. Section 3 presents the empirical framework and discusses distributional assumptions and identification. We present our estimation results in Section 4, discuss the implied substitution pattern and welfare results in Section 5, and conclude in Section 6.

2 Alpine Freight Traffic

Our data is a large-scale cross-section from the Cross-Alpine Freight Transport (CAFT) survey done every 5 years.[3] Each respective year, trucks are stopped and surveyed at all possible Alpine crossings between Vintimille and Wechsel (see Figure 2). The survey is a joint initiative by the Austrian, French, and Swiss governments to produce a representative sample of transalpine freight transport. In 2004, Germany and Italy joined the effort emphasizing the political relevance of collecting high quality data on transalpine transport activity. The data set comprises detailed information on each truck's origin and destination regions. Regions are defined at the NUTS3 level, corresponding, for example, to departments in France, districts in Germany, counties in the US. The data further include transported commodity classes, weight, vehicle characteristics, region of registration, intermediary boarder crossings, traffic direction, and more. We merge this data with shortest distances on a direct line, data on average gas and other operating costs, road and tunnel tolls, as well as with GDP data on origin and destination regions.[4] Our sample includes all French-Italian passages and all Swiss-Italian passages (see Appendix and Figure 2).

In Table 1 we present descriptive statistics of our variables. We compute *price*

[3]The available waves are 1994, 1999, and 2004 with data quality increasing in each wave.

[4]We also collected data on daily weather conditions during the sample period. Unfortunately, only the French part of the sample includes observation dates and we cannot use this information. Since mainly small and high-altitude passes are negatively affected by weather, we expect the route fixed effects to capture weather effects. To some extent, we expect the same with respect to traffic volume and congestion, even though traffic data provided by the German automobile club (ADAC) does not show significant traffic jams during the sample period.

Table 1: Alternative-specific characteristics

	Mean	Standard Deviation
Alternative characteristics		
Price	362.7	148.1
Distance	853.2	413.5
Time	18.4	11.6
User (truck) characteristics		
Weight of goods (tons)	12.8	8.8
Per capita GDP at destination	30,788.1	11,299.7

Notes: The CAFT 2004 survey includes 285,656 observations and 35,707 choice situations on French and Swiss passages.

based on an average per kilometer cost estimated by the French *Comité National Routier* on an annual basis[5] and Alpine tunnel fees. For Swiss passages, we include the Heavy Vehicle Fee which is based on ton-kilometers in Switzerland. We compute *distance* in kilometers from origin to destination regions via each respective Alpine passage using Vincenty direct geographic distance calculations implemented in Stata. We also make use of information on which border-crossings were used along the way.[6] Prato (2009) provides an up-to-date survey on the challenges in working with route choice data and emphasizes that using shortest point-to-point distances increases similarity within the choice set. Thus, we expect substitutability to be over-estimated and our consumer surplus estimates to be lower bounds. We observe that the large majority of trucks choose the alternative with the minimal price and time. This is intuitive and we therefore expect a negative impact of these variables on choice probabilities in our estimation results. We do observe, however, that a small proportion does not choose the price- and time-minimizing alternative. We can think of two main reasons for this observation. First, there may be a trade-off between time and price logistics firms face and there are some who prefer a longer route to incurring the significant tunnel fees. Second, it could be the result of unobserved heterogeneity related to logistic route choice. In order to reduce fixed costs, logistic firms may choose to combine several loads into one truck. As a consequence, some observations may not be pure Origin-Destination relationships but only the start and end points of a more complex delivery route. Furthermore,

[5]See http://www.cnr.fr/grilles_couts/e-docs/00/00/00/26/document_grille_cout.phtml.

[6]Our attempts to obtain shortest-link route distances from routing service providers such as Google Maps or Navteq failed, unfortunately. Thus, while we have a decent long-distance approximation based on manual checks on a small sub-sample of routes, we need to assume that no significant bias results from ignoring the fact that roads are not straight lines.

depending on the specific good and, for example, their service contract, we expect trucks to have (unobserved) heterogenous preferences for saving money and time. Our *time* variable is based on typical truck speed and including regular compulsory stops by European law, conditional on the number of drivers per truck. We include per capita GDP of the destination region as a user characteristic to proxy for the value of a vehicle's charge, in addition to the type of commodity. The mean in our sample is closest to per capita GDP in the Netherlands in 2004.

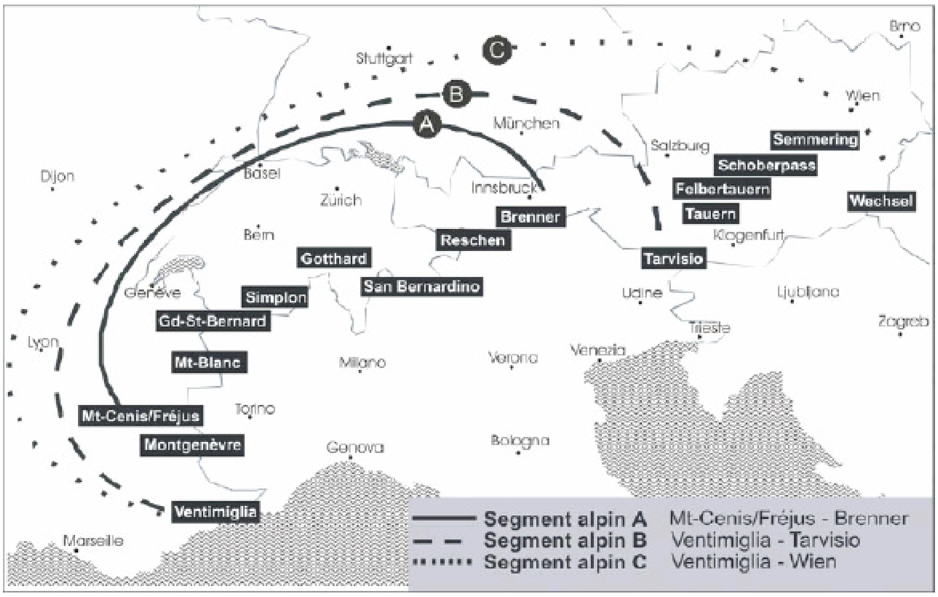

Figure 1: The Alpine corridor. Source: AlpInfo, Federal Office of Transport, Swiss Confederation.

3 Empirical Framework

We estimate the loss in user surplus from a hypothetical closure of a major transalpine road tunnel. For example, consider the problem faced by a firm located north of the Alps - say, in France - delivering its product to a downstream producer located south of the Alps - say, in Italy. By our definition, the firm faces eight mutually exclusive route options. While rail could be an option for the firm, we restrict our analysis to road freight. Even though we are forced to this restriction by our data,[7] modal choice typically depends heavily on the type of commodities

[7]For the same reason, our analysis is short-term in that we employ a static choice model and do not allow for market growth or decline through an outside option.

and logistic specificities and is, thus, largely predetermined in our choice situations. We further motivate excluding modal choice for our setup in Section 3.3.

We adopt a discrete choice framework[8] where the choice set are eight west-alpine crossings and the decision makers are individual freight trucks. In particular, user i maximizes the benefit to be obtained from a delivery trip through the Alps and faces j mutually exclusive routes. Thus, the objective function is

$$U_{ij} = \alpha_i(r_i - p_{i,j}) + x_{i,j}\beta_i + \xi_j + \epsilon_{ij}, \tag{1}$$

where r_i is the firm's revenue from the transaction, $p_{i,j}$ the price for alternative j and $x_{i,j}$ route characteristics, which both may contain interaction terms with individual characteristics, ξ_j is a constant unobserved route characteristic, and ϵ_{ij} is assumed to be independently and identically type I extreme value distributed. We have detailed user-specific information, such as GDP in destination and origin regions, commodity class, goods weight, vehicle type, location of vehicle ownership, and so on, allowing us to control for a range of observed user characteristics. However, we expect there are still user characteristics we cannot observe, such as *en route* pick-ups of goods, truck drivers' personal tastes, or special logistic needs that favor one route over another. It is common practice to model such unobserved heterogeneity by assuming parametric distributions for the relevant taste parameters. However, when estimating welfare measures, too strict assumptions will lead to biased results. Cherchi and Polak (2005) warn that assuming common mixing distributions such as the normal and log-normal may bias welfare estimations due to an inadequate representation of the true underlying distribution of tastes. Hensher and Greene (2003) give similar warnings and offer a range of simple ways for investigating sensible distributional assumptions.[9] BFKR propose a mixtures estimator that is nonparametric in the distribution of random coefficients. We have no good prior as to how our taste parameters should be distributed. We further would like to avoid assuming distributions that force portions of decision makers to have negative and unreasonably large coefficients, such as the normal distribution. Hence, we estimate both parametric and nonparametric specifications[10] and compare the welfare results.

[8]See Train (2009), McFadden and Train (2000), and Hensher and Green (2003) for in-depth treatments.

[9]One straight-forward proposition is a jack-knife approach where the researcher estimates fixed coefficient logit models on sub-samples and evaluates the distributions of estimated coefficients expected to have a non-degenerate distribution.

[10]We also coded the EM algorithm proposed by Train (2008) but encountered severe convergence problems, a common problem with EM algorithms. Fosgerau and Hess (2009) use the method of sieves to increase flexibility. They investigate the ability to recover taste distributions using two flexible approaches: In the first, they add a series expansion to a continuous base distribution using Legendre Polynomials. In the second, they employ discrete mixtures of normal distributions.

3.1 Parametric Specification

We first model user heterogeneity such that taste parameters take the following form:

$$\begin{pmatrix} \alpha_i \\ \beta_i \end{pmatrix} = \begin{pmatrix} \alpha \\ \beta \end{pmatrix} + \Pi D_i + \Sigma \nu_i$$

where

$$\nu_i \sim \mathbb{N}(0, \mathbb{I}) \tag{2}$$

and D_i is a vector of user specific characteristics in our data. User i chooses route j if and only if $U(D_i, \nu_i, p_{i,j}, x_{i,j}, \xi_j; \theta) \geq U(D_i, \nu_i, p_{i,l}, x_{i,l}, \xi_l; \theta)$ for $l = 1, ..., J$. The individual choice probability for route j follows:

$$P_{ij} = \int_\theta \frac{exp(x_{i,j}\beta_i - p_{i,j}\alpha_i + \xi_j)}{\sum_l exp(x_{i,l}\beta_i - p_{i,l}\alpha_i + \xi_l)} f(\alpha) f(\beta) d(\alpha) d(\beta) \tag{3}$$

where we assume $f(\alpha)$ and $f(\beta)$ to be normal distributions as defined in Equation 2. We estimate the parameters θ by maximum simulated likelihood. In the multinomial logit with fixed coefficients, α and β are constant across individuals so that the index i is dropped.

3.2 Nonparametric Specification

BFKR propose a general nonparametric sieve estimator of unobserved heterogeneity in a wide range of economic models. Their motivating example for the random coefficients logit model lends directly to our analysis. The idea is that the researcher has some prior over the dimensionality and range of random coefficients, that is, of unobserved heterogeneity in the utility function. Assume there are $r = 1, ..., R$ preference types in the population. We then specify a grid over the assumed support of random coefficients β. At each grid point, that is for each type r, we compute the predicted logit choice probabilities at $x_{i,j}$

$$g_j(x_{i,j}, \beta^r) = \frac{exp\left(x'_{i,j}\beta^r\right)}{\sum_{j=1}^J exp\left(x'_{i,j}\beta^r\right)}.$$

Treating these predicted probabilities as data, a simple linear regression with R predicted choice probabilities as regressors yields estimates of R weights. The left-hand side variable in this regression, $y_{i,j}$, holds the observed choices. The estimated weights are probability mass points representing the probability of observing type r in the population. Constraining probability masses to the unit interval and their sum to be equal to one, we obtain a discrete approximation to the true distribution of random coefficients. The resulting implementation is a linear inequality constrained OLS estimator.

One drawback of this simple model is the researcher's need to specify the support region where the random coefficients lie. If the latter is unknown, BFKR propose a location scale model that allows to both estimate the location of the support region and its scale. We have no good prior of the support region and thus proceed with the location scale model. A further practical advantage of the latter is the straightforward inclusion of fixed coefficients. This is an open issue in the linear estimator, given that, by definition, fixed coefficients are nonlinear parameters in the logit model. The nonlinear estimator solves the constrained least squares problem

$$\begin{aligned} \min_{a,b,\theta} \frac{1}{NJ} \sum_{i=1}^{N} \sum_{j=1}^{J} \left(y_{i,j} - \sum_{r=1}^{R} \theta^r g_j(x_{i,j}, a + b\beta^r) \right)^2 \\ \text{subject to} \quad \sum_{r=1}^{R} \theta^r = 1 \quad \text{and} \quad \theta^r \geq 0, \end{aligned} \tag{4}$$

where $a = (a_1, ..., a_K)'$ is a set of location parameters, $b = (b_1, ..., b_K)$ a set of scale parameters, θ^r the weight for the parameter vector $\beta^r = (\beta_1^r, ..., \beta_K^r)$, and

$$g_j(x_{i,j}, a + b\beta^r) = \frac{exp\left(\sum_{k=1}^{K} x_{k,i,j}(a_k + b_k\beta_k^r)\right)}{\sum_{j=1}^{J} exp\left(\sum_{k=1}^{K} x_{k,i,j}(a_k + b_k\beta_k^r)\right)}.$$

We estimate scale parameters b_K for random coefficients and set $b_K = 0$ for coefficients assumed to be fixed. We could allow all coefficients to be random but, as common with nonparametric estimators, we reach computational limits fast when increasing the number of random coefficients. Having reasonable starting values is important in nonlinear least squares estimation. We use the fixed coefficient logit estimates for the nonlinear parameters and $\frac{1}{R}$ for θ^r. Experimenting with starting values of the nonlinear parameters, we find our results are very robust to large variations in starting values. We still need to specify R grid points in the unit interval. To do so, we use the Modified Latin Hypercube Sampling method proposed by Hess et al. (2006). Compared to Halton methods, the latter has the advantage of avoiding undesired correlation patterns across dimensions while providing more uniform coverage in each dimension, and being simpler to implement.

3.3 Identification

Given their regulated nature, tolls for individual tunnels and long-distance routes vary little across time and routes. Thus, to identify demand patterns, we use individual-level variation from users' geographic dispersion across Europe. In particular, users' origin and destination locations vary relative to the locations of alpine passages. This leads to variation both in route characteristics and individual choices.

As road and tunnel tolls are not set strategically, we are confident they are not correlated with the error term. There are two factors that may affect both route choice and travel time. First, congestion may cause short-term deviation to an alternative route. On French passages, however, congestion is not a relevant problem. The Alpine Traffic Observatory, established by the European Commission and the Swiss government in 2007, finds that the Frejus and Mont Blanc tunnels rarely suffer from congestion due to heavy-duty vehicles.[11] Second, severe weather conditions may cause deviation to alternative routes while increasing travel time. Unfortunately, we cannot fully correct for this potential problem as only parts of our data have information on the date of the choice situation.[12] Once we are willing to assume that truck drivers form long-term expectations on potential obstacles for each route alternative,[13] we argue it is plausible to believe that the alternative fixed effects will capture much, if not all, potential bias.

Implicitly, we assume that we observe each individual's entire choice set and that there is no outside good. While in the long run freight expeditors may switch to other modes such as air, rail, and sea or even decide not to ship, in the short run this is very unlikely. In the first two columns of Table 4, showing market shares of alternatives before and after the Mont Blanc tunnel closure in 1999, we cannot observe a remarkable shift in market shares towards any of the three rail passages in the short-run.[14] We also do not observe a downward shift in monthly tons transported after the closure, suggesting that traffic fully deviated to alternative road passages. We interpret this as evidence that modal shift is not yet as relevant as it may be elsewhere. That interpretation is in line with anecdotal evidence on freight transport in France and Italy, citing specific logistic needs, the importance of geographical location, and freight terminals being major bottlenecks as just a few of many remaining problems preventing increased modal shift.[15] While an important issue in general, we conclude that observing only one mode is a minor drawback in our analysis.

Bajari et al. (2009) prove nonparametric identification of the distribution of random coefficients by exploiting the logit distributional assumptions on ϵ_{ij} and without relying on large support (as compared to, for example, Berry and Haile, 2010) and monotonicity restrictions. The latter makes their identification result particularly relevant for applied work. A limitation is that their proof is valid only for continuous regressors. Both regressors, for which we assume random coefficients,

[11]See, for example, the Observatory's executive summary, page 6, at http://ec.europa.eu/transport/road/doc/executive_summary_alpine_observatory_en.pdf.

[12]We did collect daily weather data for the different passages but were unable to obtain survey dates for the Swiss part of the data. This applies to congestion data analogously.

[13]Deviating from a planned route is often prohibitive in terms of cost and time.

[14]Due to missing time information at the individual level, our data do not allow direct estimation of the switch to rail and alternative roads caused by the Mont Blanc tunnel closure.

[15]See Andrews (2001), Lange and Ruffini (2007), and Peter Brett Associates LLP (2010).

are continuous.

4 Estimation results

We report results for three models of heterogeneity. These are the simple logit, the random coefficients logit with a normal distribution assumption, and the nonparametric BFKR estimator of the random coefficient distribution.[16] For conciseness, we will refer to these as Logit, RC Logit, and BFKR, respectively. In the Logit specification, observed heterogeneity in preferences can be identified by interacting individual characteristics with route characteristics. Unobserved heterogeneity is limited to the extreme value error term, which is independent of route characteristics. The Logit specification serves as an easy reference point, lending well to the investigation of a variety of utility specifications, and provides reasonable starting values for the RC Logit and BFKR models. In the latter, we report two specifications with one and two random coefficients. In the RC Logit, we assume random coefficients to be normally distributed, which is by far the most common distribution assumption in the literature employing random coefficients logit models. In the BFKR model, we make no assumption on the form of preference distribution, whatsoever, but assume that our discrete approximation is a valid representation of the true distribution.

Tables 2 reports estimations results with one random coefficient, Table 3 reports results with two. Both tables show that the estimated coefficients have signs as expected. In particular, price and time have negative signs. The coefficients of the interaction terms between weight of goods transported and the alternative specific constants show that heavily loaded trucks are less likely to use Swiss passages and the Montgenevre passage.[17] This corresponds to our expectation that the Montgenevre pass, having the highest elevation, is less attractive to heavy vehicles likely due to increased fuel consumption and safety concerns on steep slopes. In Switzerland, extra incentives are given for transiting heavy goods vehicles to switch to rail on their passage through the Alps. While we account for monetary incentives by including the extra fees in our price variable, the weight interaction terms may capture further incentives we cannot observe in our data.

In Table 2, price coefficients are higher, in absolute terms, in the RC Logit and BFKR specifications than in the Logit. Hence, not modeling unobserved heterogeneity not only fails to account for the spread of preferences but also biases the

[16]We run our parametric estimations using the *mixlogit* Stata command by Hole (2007). For nonparametric estimations, inference, and compensating variation, Matlab code is available from the authors on request.

[17]The base category both for the alternative specific constants and for the interaction term with good weights is alternative 5, the Mediterranean passage at Vintimille.

Table 2: Results with one random coefficient (1)

	Logit	RC Logit Mean	RC Logit SD	BFKR
Price	-8.478***	-9.298***	1.523***	*See Figure 4*
	(.118)	(.149)	(.079)	
Time	-.922***	-.869***		-.854***
	(.230)	(.232)		(.040)
Price × $GDP_{destination}$	1.153***	1.533***		1.096***
	(.294)	(.317)		(.001)
Mont Blanc	4.214***	4.291***		3.969***
	(.280)	(.297)		(.243)
× Weight	.008	.010		.014***
	(.006)	(.006)		(.001)
Frejus	5.144***	5.256***		4.488***
	(.234)	(.254)		(.318)
× Weight	.020***	.023***		.026***
	(.005)	(.006)		(.002)
Montgenevre	-5.843***	-6.657***		-4.886***
	(.552)	(.657)		(.291)
× Weight	-.031***	-.036***		-.024***
	(.009)	(.010)		(.005)
Gd St-Bernard	.154	.178		.632***
	(.310)	(.332)		(.065)
× Weight	-.058***	-.063***		-.072***
	(.009)	(.010)		(.004)
Simplon	-.095	-.113		.428***
	(.310)	(.330)		(.089)
× Weight	-.047***	-.051***		-.085***
	(.010)	(.010)		(.004)
St. Gotthard	3.576***	3.730***		4.128***
	(.293)	(.312)		(.196)
× Weight	-.073***	-.076***		-.096***
	(.007)	(.007)		(.007)
San Bernadino	2.654***	2.789***		3.282***
	(.297)	(.316)		(.102)
× Weight	-.073***	-.075***		-.090***
	(.008)	(.008)		(.005)
Likelihood ratio	105342.50	130.60		
Prob > χ^2	.000	.000		
Pseudo R^2	.71			

Notes: All specifications include route dummy-commodity class and time-commodity class interaction terms as well as route dummies interacted with dummies indicating traffic connecting Italy with regions west and north of the Alps, respectively. The reference route is the Mediterranean crossing at Vintimille. Standard errors are reported in parenthesis, choice situation-clustered robust standard errors in BFKR estimation. 500 Halton draws used for simulations in parametric random coefficients logit estimations. 285,656 observations.
*** p< 0.01, ** p< 0.05, * p < 0.1

estimated means. The means and standard deviations of the price coefficient are significant in both random coefficient specifications. Figure 4 shows a decent Gaussian shape, confirming our parametric distribution assumption in the RC Logit. Hence, we can be reasonably confident that unobserved heterogeneity with respect to price plays an important role in our freight transport setting. In Table 3, we allow individuals to have heterogenous (and potentially correlated) preferences over both price and time. Interpreting the RC Logit result, where the standard deviation of the time coefficient is not statistically different from zero, we conclude the estimated distribution to be degenerate. In economic terms, we may be tempted to conclude that there is no unobserved heterogeneity in preferences over time. Observing the BFKR estimates in Figure 4, however, we clearly see significant heterogeneity both in preferences over price and time. A priori, it is not obvious which distribution to assume. It is difficult to proceed in an ad-hoc fashion by assuming various readily available parametric distributions and using the ones yielding significant parameters estimates. There is no clear rule which and how many parametric distributions to investigate before 'giving up' and coming to the conclusion that here may be no unobserved heterogeneity in the data.

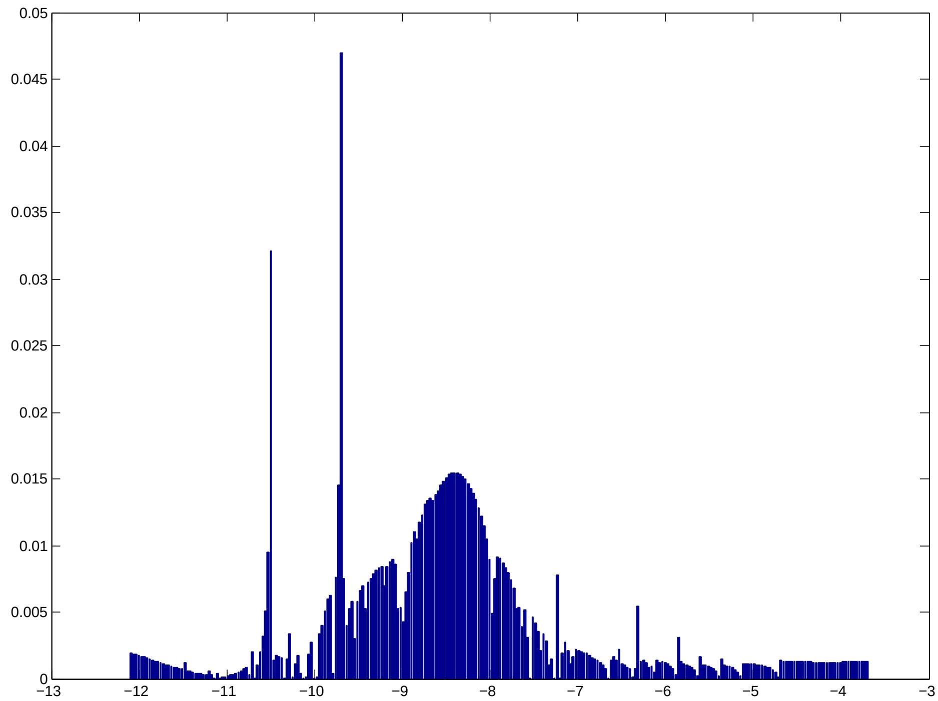

Figure 2: Distribution of price coefficient from BFKR estimation in Table 2. Regimes R=274.

The main assumption in BFKR is that their sieve estimator is a discrete approximation of the true distribution. Therefore, the number of grid points is a key

Table 3: Results with two correlated random coefficients (2)

	Logit	RC Logit Mean	RC Logit SD	BFKR
Price	-8.478***	-9.244***	1.273***	*See Figure 4*
	(.118)	(.148)	(.078)	
Time	-.922***	-1.504***	.271	
	(.230)	(.266)	(.291)	
Price × $GDP_{destination}$	1.153***	1.520***		1.160***
	(.294)	(.322)		(.001)
Mont Blanc	4.214***	4.276***		5.913***
	(.280)	(.293)		(.316)
× Weight	.008	.010*		.018***
	(.006)	(.006)		(.002)
Frejus	5.144***	5.254***		6.753***
	(.234)	(.248)		(.403)
× Weight	.020***	.023***		.032***
	(.005)	(.005)		(.003)
Montgenevre	-5.843***	-6.638***		-5.318***
	(.552)	(.665)		(.319)
× Weight	-.031***	-.035***		-.013***
	(.009)	(.010)		(.003)
Gd St-Bernard	.154	.237		1.553***
	(.310)	(.332)		(.067)
× Weight	-.058***	-.065***		-.071***
	(.009)	(.010)		(.004)
Simplon	-.095	-.064		1.485***
	(.310)	(.330)		(.089)
× Weight	-.047***	-.053***		-.092***
	(.010)	(.011)		(.004)
St. Gotthard	3.576***	3.818***		5.434***
	(.293)	(.310)		(.225)
× Weight	-.073***	-.077***		-.109***
	(.007)	(.007)		(.008)
San Bernadino	2.654***	2.872***		4.509***
	(.297)	(.314)		(.115)
× Weight	-.073***	-.076***		-.102***
	(.008)	(.008)		(.006)
Likelihood ratio	105342.50	262.77		
Prob > χ^2	.000	.000		
Pseudo R^2	.71			

Notes: All specifications include route dummy-commodity class and time-commodity class interaction terms as well as route dummies interacted with dummies indicating traffic connecting Italy with regions west and north of the Alps, respectively. The reference route is the Mediterranean crossing at Vintimille. Standard errors are reported in parenthesis, choice situation-clustered robust standard errors in BFKR estimation. 500 Halton draws used for simulations in parametric random coefficients logit estimations. 285,656 observations.
*** p< 0.01, ** p< 0.05, * p < 0.1

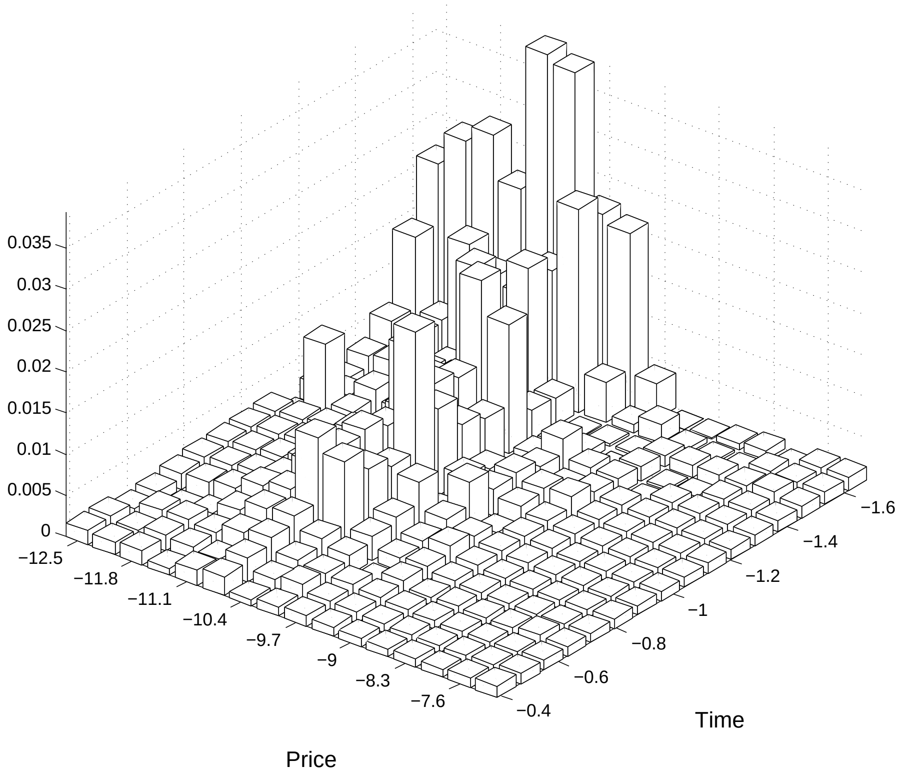

Figure 3: Distribution of price and time coefficients from BFKR estimation in Table 3. Regimes R=256.

parameter the researcher needs to define. It is limited by sample size and computer memory. With 8GB RAM and our sample size of 35,707 choice situations, we are able to set R=274 with one and R=256 with two random coefficients, that is $R = \frac{N}{130}$ whereas BFKR propose $R = \frac{N}{40}$ in their Monte Carlo experiment. Furthermore, to use parametric methods for computing standard errors of the BFKR estimates, we need to assume that the grid points r are the true types that generated the data. It is obvious that, unless we can specify an almost infinite number of r, estimated probability mass points and thus their standard errors can only be approximations to the truth. We compute 95% confidence intervals, following BFKR and Gallant (1975), using standard errors from unconstrained nonlinear regression and clustering by individual choice situations. Confidence intervals are then defined as $CI_{95} = \left[\hat{\theta}^r - 1.96 \cdot SE\left(\hat{\theta}^r_{unc}\right), \hat{\theta}^r + 1.96 \cdot SE\left(\hat{\theta}^r_{unc}\right)\right] \cap [0,1]$, where $SE = \sqrt{\hat{\nu}}$ and

$$\hat{\nu} = \frac{NP}{NP-1}\left(\frac{SSE}{N-p}(J'J)^{-1}\left(\sum_{k=1}^{M}\left(\sum_{j\in G_k} u_j' \sum_{j\in G_k} u_j\right)\right)\frac{SSE}{N-p}(J'J)^{-1}\right).$$

However, as do BFKR, we find these confidence intervals to be very conservative and, thus, too uninformative to report. From our experience with estimations using varying grids sizes, this problem seems to get worse as the number of grid points R increases and θ are estimated closer to the zero boundary. Bootstrapping is infeasible as the nonlinear least squares routine consumes significant computing time with the size of our sample and parameter vector.

5 Counterfactual Analysis

5.1 Diversion pattern

To verify the structural quality of our estimates, we compare counterfactual shares with those observed after the Mont Blanc tunnel closure in 1999. Road tolls and tunnel fees were not adjusted as a consequence of the closure and there is currently no road congestion pricing on the relevant road network. Thus, we assume there is no strategic pricing and prices remain fixed after a hypothetical future closure of a tunnel. In Table 4, we present the substitution due to a hypothetical closure of the Mont Blanc tunnel in 2004 predicted by the BFKR model with two random coefficients. We compute market shares in the table based on tons transported.

Table 4: Passage Market Shares (%) - Counterfactual closure in 2004

		1999		2004	
		Open	Closed	Open	Closed
Road	Mont Blanc	21.79		10.14	
	Frejus	27.68	44.54	32.51	35.87
	Montgenevre	1.09	3.38	0.67	2.31
	Vintimille	21.26	24.36	33.76	33.31
	Gd St-Bernard			1.19	1.99
	Simplon			1.25	1.54
	St. Gotthard			18.07	21.59
	San Bernadino			2.41	3.38
Rail	Basel (CH)	10.61	10.84		
	Modane (Frejus)	15.42	15.21		
	Vintimille	2.16	1.67		
Monthly tonnage (1000s)		4,455	4,619	4,038	

Notes: 2004 Predictions are simulated using specification (2) of the BFKR model. We use 1000 quasi-random draws to compute choice probabilities.

Table 4 shows that most traffic diverted to the nearby Frejus tunnel when the Mont Blanc option dropped out in 1999. The traffic share at the Mont Blanc dropped from 21.79% to 0%, causing an increase of the traffic share at the Frejus from 27.68% to 44.54%. We do not have monthly data for Swiss passages in 1999, limiting comparison to the French passages. A further limitation arises from the fact that, between 1999 and 2004, Switzerland gradually increased weight restrictions for heavy duty transit. For a large share of trucks, Swiss passages were not an available option in 1999 while, in 2004, these limitations were virtually gone. Since 1999, freight traffic through the Mont Blanc tunnel has been reduced significantly, its share being 10.14% in 2004. Our counterfactual results show the largest shifts from the Mont Blanc tunnel to the Frejus and St. Gotthard passages. While 10.14% traffic share need to be compensated by the remaining passages, the traffic share at the Frejus is predicted to increase from 32.51% to 35.87% and at the St. Gotthard from 18.07% to 21.59%. We interpret the similar tendencies on French passages throughout 1999 and 2004 to confirm the fit of the BFKR model but stress the limited comparability of these market shares due to institutional and observational differences.

5.2 Consumer surplus

We analyze a hypothetical closure of the Mont Blanc tunnel in 2004 and compute users' compensating variations, that is, the amount of money one would have to give to infrastructure users to maintain their ex ante utility levels. We define the ex ante situation as the reference point, as suggested by Trajtenberg (1989). In the multinomial logit model, abstracting from observed heterogeneity for notational simplicity, computation of the compensating variation is straight forward and given by the difference of the ex post and ex ante values of the logsum measure with no unobserved taste heterogeneity:

$$CV = \frac{1}{\alpha}\left\{\ln\sum_j \exp\left(\beta' x_j^{pre}\right) - \ln\sum_j \exp\left(\beta' x_j^{post}\right)\right\} \tag{5}$$

As the random coefficients logit model introduces unobserved taste heterogeneity, each individual now may have her own valuation of route characteristics. We integrate over the estimated mixing distributions by simulation and compute the mean and total compensating variation. This is where the *how* we model heterogeneity comes in. With random coefficients, any distributional assumption has a direct impact on the consumer surplus measure which, following Train (1998) and von Haefen (2003), we compute as:

$$CV_i = \int \frac{1}{\alpha_i}\left\{\ln\sum_j \exp\left(\beta' x_{i,j}^{pre}\right) - \ln\sum_j \exp\left(\beta' x_{i,j}^{post}\right)\right\} f\left(\alpha \mid \theta^{pre}\right) d(\alpha) \tag{6}$$

Equations 5 and 6 imply the assumption that the marginal utility of income, α_i, is independent of income. That is, indirect utility is additive and linear in income. While this is a restrictive assumption, Train (2009) points out, on page 57, that it needs to hold only ‘over the range of implicit income changes that are considered by the policy’. This means that if individual compensating variations are low relative to income, which is arguably true for our case, the assumption does not need to hold in general but only for the considered small range. Maintaining this assumption significantly simplifies our computations. We solve Equation 6 for each individual via simulation by sampling from the estimated mixing distributions. To obtain the population mean and total change in user benefits, we weight the estimated sample means of compensating variations using the expansion factor in the CAFT 2004 data.

Table 5: Welfare effects of tunnel closure

Compensating variation (in 2004 Euros)	Logit	RC Logit		BFKR	
		(1)	(2)	(1)	(2)
Sample mean	1.11	.84	.69	1.80	1.39
Population mean	1.28	.96	.79	1.88	1.43
Population total (Mio)	4.83	3.62	2.97	7.09	5.39

Notes: We use 1000 quasi-random draws to solve the integral in Equation 6.

In Table 5, we report the compensating variation for closing the Mont Blanc tunnel as the unweighted and weighted means over the sampled individuals, and as the weighted sum yielding the population total. The BFKR estimates imply economically significantly higher losses in user surplus. With two random coefficients, the BFKR estimate implies a loss of €5.39 Mio and the RC Logit a loss of €2.97 Mio. With one random coefficient, the BFKR estimate is almost double that of the RC Logit estimate, €7.09 Mio versus €3.62 Mio. Compared to both the Logit and the BFKR estimates, both RC Logit specifications underestimate the loss in consumer surplus. We admit that a more informative comparison would include confidence intervals for these estimates. These can be computed using the delta method. However, they rely on the estimated variances of $\hat{\theta}^r$. Since there is currently no method to estimate correct confidence intervals, for the reasons elaborated in Section 4, we are not able to provide these inference results for our welfare estimates. The large relative differences between the RC Logit and BFKR results strongly suggest, however, that modeling unobserved heterogeneity in a more flexible way can lead to severely differing economic conclusions.

6 Conclusion

Estimating welfare implications of specific policy measures is relevant in many economic applications. The discrete choice framework provides a convenient way to estimate the compensating variation, for example for changes in choice sets or in product characteristics, when individuals face a set of mutually exclusive choices. It is well known, that modeling heterogeneity is key to understanding preferences and hence to quantify consumer surplus. While the random coefficients logit model offers a highly flexible way to approximate any random utility model arbitrarily well, it hinges on the assumption that the researcher knows the correct distribution of random coefficients a priori. While estimating welfare effects is a common exercise in many applications, we are not aware of many studies exploring the role of unobserved heterogeneity in this context. We provide insights into the importance of *how* unobserved heterogeneity is modeled when estimating welfare effects of policy measures. We do so by comparing consumer surplus estimates from a recently proposed nonparametric estimator of preference distributions and the standard parametric random coefficients logit model. To our knowledge, we are the first to apply the BFKR estimator to real-world data in a random coefficients logit setting. Employing revealed preference data, we analyze the implications of a much debated transport policy measure: the closure of the alpine Mont Blanc tunnel to freight traffic.

We estimate the annual loss in user benefits ranges from 2.97 to 3.62 million Euros in the RC Logit model, while our BFKR estimates imply annual losses ranging from 5.39 to 7.09 million Euros. Hence, in our analysis both parametric assumptions and the dimensionality of modeled unobserved heterogeneity have a significant impact on welfare results. We thus caution the exclusive use of standard distributional assumptions in modeling heterogeneity and demonstrate the simple implementation of the BFKR estimator.

Acknowledgements

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Appendix: Alpine passages

Our sample includes the following passages:

1. Mont Blanc tunnel
2. Frejus tunnel
3. Montgenevre pass
4. Vintimille expressway (along the Mediterranean coast)
5. Grand St-Bernard tunnel
6. Simplon pass
7. St. Gotthard tunnel
8. San Bernadino pass

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Chapter 3

Effort in Nomination Contests: Evidence from Professional Soccer

joint with Jeanine Miklós-Thal

1 Introduction

Situations abound in which several candidates compete for a limited number of desirable positions and selection is based on the candidates' relative reputations. Employees compete for promotions, given to the employee who their superior believes will be most effective in the higher-level position. Hiring decisions are based on subjective comparisons of candidates' skills and potentials. Political parties nominate election candidates on the basis of their anticipated abilities to attract voters. Team coaches in sports select those players for important matches who they believe will lead their teams to victory.

While motivating employees is often an explicit goal of promotion systems, the decision-maker's objective in a hiring contest is usually simply to select the most able agent.[1] Irrespective of a contest's ultimate goal, however, contest participation can have important incentive effects. Whenever current performance affects perceived ability, and thereby potentially also the contest outcome, actions aimed at improving one's performance can be profitable.

Does contest participation always motivate agents, and, when it does, what determines the extent of the effect? We propose a simple theoretical model predicting

[1]Prendergast (1999) provides an excellent survey of incentive provision in firms. Chan (1996) analyzes the conflict between motivating internal agents by the prospect of a promotion and selecting the most promising candidate out of a pool of internal and external candidates.

that each candidate's effort incentive depends on his own and his rivals' current reputations. Candidates who have realistic chances of being selected but are not too confident have strong incentives to exert higher than normal effort. Candidates in very weak or very strong positions, on the other hand, do not have much to gain from exerting additional effort, since changes in their performances are unlikely to affect the final decision. In some contexts, higher effort also increases the risk of an injury or leads to exhaustion. When competing for a position that requires continued fitness, candidates who are confident their reputations sufficiently exceed those of other contestants may therefore find it optimal to exert less than normal effort.

We use readily available data from professional soccer to test these predictions. When a nation qualifies for an international tournament, such as the Soccer Euro Cup, the national team coach gets charged with nominating a fixed number of players for the Cup.[2] Nationality determines the set of legally eligible players and hence whether a player participates in the nomination contest for a specific national team. A Euro Cup participation is clearly a milestone in any player's career.

The key feature of professional soccer that allows us to estimate the effects of nomination contests is the coexistence of important tournaments between national teams with international player compositions of club teams. We use a panel data set of all players that worked for clubs in the German Soccer League (*1. Bundesliga*) in the seasons 2006/07 and 2007/08. About two thirds of the players belong to nations that took part in the Euro Cup, the most important international soccer Cup alongside the World Cup, in summer 2008.[3] This set of players will provide the treatment group in our empirical analyses. In players from nations that did not participate in the so-called *Euro 2008* we have an exceptionally good control group, since these players work in exactly the same environment as players from qualified nations but did not face the additional career opportunity of the upcoming Euro Cup. The treatment period starts on the day a player's nation qualified for the *Euro 2008*.

Our data contain individual performance measures of two types. First, individual outputs such as shots on goal, ball contacts, passes received, and the number of minutes played. Second, performance grades assigned to players by soccer magazines after each match.

To distinguish between players with different nomination chances, we construct a time-varying variable that measures how frequently a player was selected for his national team in the more recent past. Difference-in-difference-in-differences analyses show that for players with intermediate chances the Euro Cup qualification

[2]National team compositions are flexible in friendly matches between nations or qualification matches for international Cups, but not in international Cups.

[3]The Euro Cup and the World Cup take place every four years, and are always two years apart from each other. There are some other international cups, such as the Copa America or the Africa Cup of Nations, but these are far from being as important (in terms of media coverage, premia paid by national teams, etc.) as the Euro and the World Cup.

treatment had a positive impact on many performance measures. For instance, the estimated increase in the number of passes such players receive per minute is 11%. The empirical results also confirm that injury and exhaustion concerns matter: for players with very high nomination chances, the impact of nomination contest participation is negative across a variety of output measures. Moreover, for duels, which carry a particularly high injury risk, all statistically significant effects are negative. Consistent with the theory, we find no impact on the performances of players without past national team appearances.

For players with intermediate chances, our study hence confirms that " ... the increased rivalry benefits clubs, because players exert even higher effort in their clubs in order to get into the national team.", as claimed by Oliver Bierhoff, manager of the German national team (Handelsblatt, 9/4/2009).[4] An upcoming Cup can be to the detriment of clubs that employ regular players of national teams, who are highly certain of their nominations, however. One may only speculate that statements such as "We want to ignite rivalry, and we want it for every position." (stern.de, 11/8/2004) by the German national team coach Joachim Löw are meant to reassure clubs in this respect.[5]

Related literature We are not aware of any other empirical study of nomination contests. There is however a sizeable literature on rank-order tournaments, in which agents' outputs during the tournament fully determine payoffs. An agent who starts out as a favorite still needs to outperform all his rivals to win, while an underdog does not face any handicap.[6] Many if not most hiring and promotion decisions are instead based on relative reputations, that is, on assessments of agents' relative abilities that incorporate not only recent but also past achievements and other relevant information. In the nomination contests for soccer teams, for example, two players who perform equally well during the nomination period will not be nominated with equal probabilities if one of them starts out with a higher reputation than the other.

The literature on rank-order tournaments is related to our paper because the predicted relation between an agent's winning probability and his effort incentives is similar. Most empirical studies of rank-order tournaments, however, focus on the more basic question whether higher prize differentials lead to more effort. Ehrenberg and Bognanno (1990) and Orszag (1994) provide evidence from golf tournaments, Becker and Huselid (1992) look at auto racing, and Knoeber and Thurman (1995) examine the impact of tournament-style contracts in the broiler industry. Garicano

[4]The original quote in German is "... der größer werdende Konkurrenzkampf bereichert auch die Vereine, weil die Spieler sich in ihren Klubs noch mehr anstrengen, um in die Nationalmannschaft zu kommen" (Handelsblatt, 9/4/2009).

[5]The original quote in German is "Wir wollen den Konkurrenzkampf entfachen, wir wollen ihn auf jeder Position haben." (stern.de, 11/8/2004).

[6]A special case are biased tournaments (Meyer 1991, 1992) in which contestants face different handicaps. Biased tournaments are theoretically equivalent to contests based on relative reputations in a special case only. See footnote 11 in section 2 for more details.

and Palacios-Huerta (2006) show that higher prize differentials increase not only creative but also destructive effort (in the form of fouls) in soccer.[7]

More closely related to our paper, Brown (2010) shows that superstar Tiger Woods' participation in golf tournaments adversely affects the performances of his rivals. The impact is particularly strong for (higher skill) exempt players who would have realistic winning chances in the absence of Woods. Our study differs along several dimensions (in addition to looking at nomination contests instead of tournaments). By constructing a variable that measures players' relative national team nomination chances, we can test predictions about the impact of contest participation for players with winning chances from zero to virtually one. Brown (2010) instead compares situations - without and with Tiger Woods - in which other exempt players have either intermediate or low winning chances. Moreover, the institutional characteristic that players of many different nationalities work for the same clubs but only some nations participate in the Euro Cup allows us to test for causal effects of contest participation,[8] whereas Brown (2010) and other empirical studies compare tournaments with different features.[9]

Our motivating theory incorporates signal jamming, as in Holmström's (1982) seminal paper on career concerns, into the classic rank-order tournament model of Lazear and Rosen (1981).[10] Höffler and Sliwka (2003) use a similar theory to study the potential benefits of managerial turnover in revitalizing rivalry between employees. We propose a model that is closer to the nomination contests in our empirical application and focus on the equilibrium relation between individual effort and winning chances instead. Relative reputational concerns have also been studied in theoretical models on rivalry between experts (Effinger and Polborn 2001, Ottaviani and Sorensen 2006).

The next section develops a theory of nomination contests and derives empirical predictions. Section 3 describes the data, our choice of output measures, and the institutional context. Section 4 explains and discusses the empirical strategy. Section 5 contains the empirical results. Section 6 offers a brief conclusion and implications for other situations.

[7] Similar in spirit, Duggan and Levitt (2002) find that there is more corruption in sumo matches in which one wrestler faces a particularly high marginal payoff from winning.

[8] Miguel, Saiegh and Satyanath (2008) exploit international compositions of soccer teams to test whether there is a connection between cultural background and violence on the field.

[9] Another recent related study is Franke (2010) who shows that amateur golfers perform better in tournaments where individual scores are evaluated relative to a player's handicap than in standard tournaments. Sunde (2009) finds a negative correlation between the heterogeneity of opponents and the number of games in tennis matches.

[10] On the theory of rank-order tournaments, see also Green and Stockey (1983), Dixit (1987), Meyer (1992), Baik (1994), Moldovanu and Sela (2001), and the above-mentioned survey by Prendergast (1999).

2 Theory

Suppose there are two agents (for example, two soccer players of the same nationality), one of whom can be selected for an attractive post at the end of a fixed time period. The nomination decision is taken by a principal (the national team coach) whose objective is to select the most skillful agent. Hence, unlike in a classic rank-order tournament à la Lazear and Rosen (1981), it is the principal's beliefs about the agents' skills that determine the winner.

We model learning about each individual agent's skill as in Holmström (1982). Let η_j denote agent j's ($j \in \{1,2\}$) skill level, which is assumed to be constant over the relevant time period. At the beginning of the nomination contest, the agents and the principal share the same prior beliefs. Specifically, we assume that the prior of η_j follows a normal distribution with mean m_j and precision (equal to the inverse of the variance) $h_j > 0$. The prior distributions of η_1 and η_2 are independent. Over time, learning about η_j occurs through the observation of j's performance. For simplicity, we consider learning in a single time period, called the nomination period. Agent j's output in the nomination period is given by

$$y_j = \eta_j + a_j + \varepsilon_j,$$

where $a_j \in [0, \infty)$ is j's effort in the nomination period, unobservable for the principal and agent $k \neq j$. ε_j is a stochastic noise term, and we assume that ε_1 and ε_2 are independently and normally distributed with zero means and precision $h_\varepsilon > 0$.

In addition, each agent faces an injury risk, modelled as an increasing function $r(\cdot)$ of individual effort with $r(0) \geq 0$ and $\lim_{a \to \infty} r(a) \leq 1$. The principal's objective is to nominate the most skillful agent, conditional on that agent not being injured. If both agents remain injury-free, then after observing y_1 and y_2 the principal will select $j \neq k$ whenever[11]

$$E[\eta_j \mid y_j] > E[\eta_k \mid y_k]. \tag{1}$$

If exactly one of the agents is injured, the principal will select the other agent. If both agents are injured, none will be selected.

The expected payoff of agent $j \neq k \in \{1,2\}$ is

$$\begin{aligned}&(1 - r(a_j))(1 - r(a_k)) \Pr\{E[\eta_j \mid y_j] > E[\eta_k \mid y_k]\} W_j \\ &+ (1 - r(a_j)) r(a_k) W_j + S_j(a_j) - c_j(a_j),\end{aligned}$$

where $W_j > 0$ denotes the (expected) prize j receives if the principal selects him. The function $S_j(a_j)$ measures agent j's expected gross payoff in the absence of the

[11] If $h_1 = h_2$, then there exists a biased rank-order tournament as in Meyer (1991, 1992) that is equivalent to the decision rule in (1). In a biased tournament, the contestant with the lower prior reputation has to outperform the other agents by a given amount to win. For $h_1 \neq h_2$, the rates at which the principal updates his beliefs about the agents' skills as a function of observed outputs differ, and therefore there is no direct equivalence with a biased tournament.

nomination contest and $c_j(a_j)$ his disutility of effort. We assume that $S_j(a_j) - c_j(a_j)$ is strictly concave and reaches a unique maximum at

$$a_j^n > 0,$$

the "normal" effort level of player $j \in \{1, 2\}$.

In a Bayesian Nash equilibrium, each agent's effort choice must be optimal given the other agent's effort choice and beliefs, and the principal must correctly anticipate effort choices. Appendix A contains a detailed analysis of the equilibrium conditions and comparative statics with respect to the equilibrium effort levels (a_1^*, a_2^*).

The main results are as follows. First, in the benchmark case without any injury concerns (i.e., $r(a) = 0$ for all a) we always have $a_j^* > a_j^n$. In this case, a_j^* depends on $\Delta = |m_1 - m_2|$ but not on m_1 and m_2 individually, and

$$\frac{da_j^*}{d\Delta} < 0 \text{ if } \Delta > 0,$$
$$\frac{da_j^*}{d\Delta} = 0 \text{ if } \Delta = 0.$$

As m_j varies, the relation between j's equilibrium effort and equilibrium winning probability is a symmetric inverted U-shape with a maximum at winning probability 50%. As j's equilibrium winning probability approaches 0 or 1, respectively, a_j^* goes to a_j^n.

If the injury risk function is increasing instead, the effort impact of the nomination contest is ambiguous. Intuitively, $a_j^* < a_j^n$ when the marginal effect of higher effort on j's winning probability is small but j has a good winning chance conditional on remaining injury-free. Ceteris paribus, this is the case if m_j is sufficiently high so that j's winning probability is close enough to 1 but the marginal effect of effort on the winning probability is close to 0. If on the contrary agent j has a very low winning chance, the contest will not affect his effort significantly: $\lim_{(m_j - m_k) \to -\infty} a_j^* = a_j^n$. For intermediate winning chances and sufficient uncertainty about the agent's ability, the winning concern dominates the injury concern ($a_j^* > a_j^n$) as long as the injury risk function is not too steep. However, a_j^* as a function of the equilibrium winning probability always reaches its maximum at a winning probability strictly below 50% now. Figure 1 depicts the relation between agent 1's equilibrium winning probability and his equilibrium effort as his prior reputation m_1 varies in a numerical example. The horizontal line indicates the normal effort level a_1^n the player would exert in the absence of the nomination contest. The equilibrium effort is increasing in the agent's equilibrium winning probability for low winning chances, but decreasing for higher winning chances. Moreover, because of the injury risk the equilibrium effort is maximal at a winning chance below 0.5, and lies below a_1^n if agent 1 has an equilibrium winning probability sufficiently close to 1.

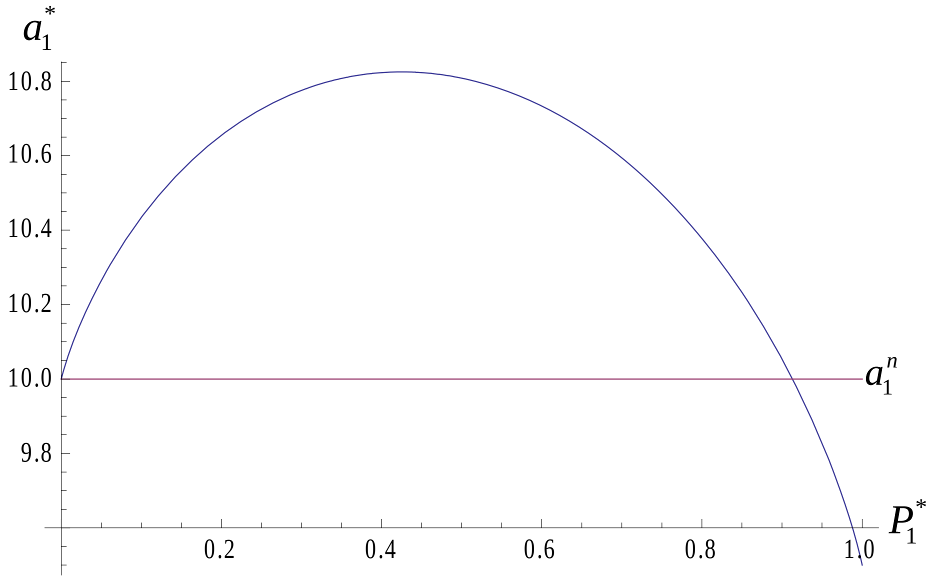

Figure 1: Equilibrium relation between agent 1's effort level a_1^* and his winning probability P_1^*. $W_1 = 10$, $m_2 = 1$, $h_1 = h_2 = 2$, $h_\varepsilon = 1$, $S_1(a) = S_2(a) = 10a$, $c_1(a) = c_2(a) = \frac{a^2}{2}$, $r(a) = 0.05a$ for $a < 20$ and $r(a) = 1$ for $a \geq 20$.

In summary, the theory predicts that nomination contest participation leads to higher than normal effort if an agent has realistic winning chances but is not too certain of winning either. For agents with very good winning chances, the prediction is that nomination contest participation leads to less than normal effort as long as injury concerns are relevant. In the empirical analysis, we will study the evolutions of observable output and performance measures to test these predictions. The interpretation is that changes in effort (training intensity, motivation and concentration on the field, lifestyle, ...) lead to changes in performance and can hence be detected by looking at performance.

3 Institutional Characteristics and Data

3.1 *Euro 2008* qualifications and national team nominations

Our empirical analyses focus on the time period between the end of the World Cup 2006 on July 9, 2006, and the end of the 2007/08 soccer season on May 17, 2008. The *Euro 2008* began on June 7, 2008. As illustrated in the timeline in Figure 2, the qualification matches for the *Euro 2008* started shortly after the World

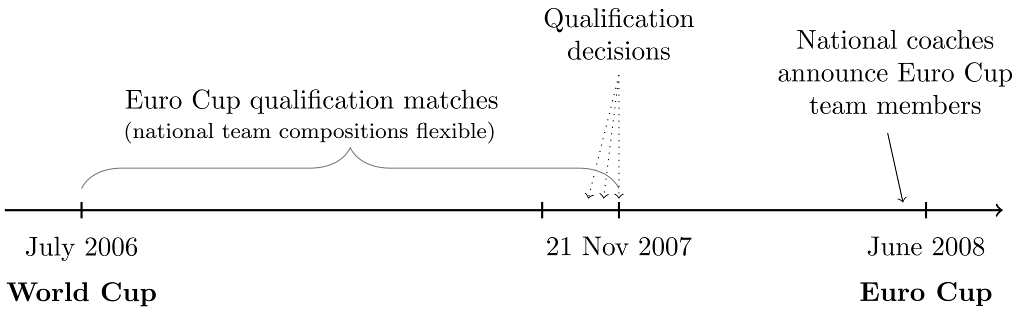

Figure 2: Timeline

Cup. All eligible nations, fifty in total for the *Euro 2008*, usually participate in the qualification matches. The official announcement of qualified nations took place on November 21, 2007, but several nations de facto qualified before that date having won sufficiently many matches. A group of four countries (Czech Republic, Germany, Greece, and Romania) qualified about one month before the official date, on either the 13th or 17th of October, while ten other nations qualified on the 17th or 21st of November. The two remaining participants were Austria and Switzerland, the host nations, which by the rules of the Cup participate automatically. We exclude players with citizenship of these two countries from all the empirical analyses.

National coaches can select different players for every non-Cup national team match if they wish to do so, and as we will document there is indeed considerable temporal variation in national team compositions for non-Cup matches. For the *Euro 2008*, however, all coaches had to nominate a fixed selection of 23 players. The deadline for the coaches' announcements of their team selections was May 28, 2008, eleven days after the end of the German soccer season. There were some differences between qualified countries regarding the date and procedures according to which national coaches announced their decisions, but most coaches made their final statements either between the last but one and the last, or after the last game day of the German soccer season.

A number of other international tournaments took place in the relevant time period: the Copa America in July 2007, the Africa Cup of Nations in January 2008, and the 2008 Olympic summer games in August 2008. These Cups could potentially interfere with our analysis by creating similar incentives as the *Euro 2008* but for different groups of players. However, because of their limited media coverage and endorsement opportunities, participation in these international tournaments is considerably less attractive for players than a Euro (or World) Cup participation. Some clubs do not even allow their players to miss club activities in order to participate.[12] Formally testing for an incentive effect of the Copa America, using the

[12]For example, Bundesliga clubs Schalke 04 and Werder Bremen clashed with the Brazilian

same empirical strategy as described below for the *Euro 2008*, we found no evidence of any effect. We therefore feel that it is safe to ignore other international Cups for the purpose of this paper.

3.2 Data and output measurement

We use a panel data set that contains detailed player-game day level information about the German Soccer League (*1. Bundesliga*) in the seasons 2006/07 and 2007/08.[13] The data provide individual output measures for all participating players in each match. In addition, we constructed a panel data set of the performance grades that two major German soccer magazines, *Kicker* and *Sportal*, assign to players after each match. We matched these data sets with data about individual injuries collected by a firm that runs an online fantasy soccer game.[14] Finally, we collected data on all national team participations of players in our sample between summer 2005 and the *Euro 2008* using publicly available sources.[15]

Our unit of observation is a player-game day.[16] In the analyses herein, we restrict attention to players for whom we have observations both before and after the official *Euro 2008* qualification date (November 21, 2007), and who were on the field at least once in the 2006/07 season as well as in the 2007/2008 season. We also exclude goalkeepers, because they have very different tasks than field players and many of our output measures are not applicable to them. The remaining number of observations is 11, 799, including observations where a player spends the entire time on the reserve bench. There are 18 teams in the *Bundesliga* and 216 matches per season.

Table 1 lists the nationalities of the players in our sample. The treatment group consists of all players whose nations participated in the *Euro 2008*. Players of all other nationalities are in the control group. About half the players are German, while the rest originate from all over the world. The *Bundesliga* was the best represented national League in the *Euro 2008*, with active players in fourteen out of sixteen national teams.

The *Bundesliga* data contain a variety of detailed individual output measures:

Shots on goal - The ultimate objective in soccer is to shoot goals and prevent goals by the opponent. Shots on goal includes actual goals, but also failed goal attempts. The main advantage of using shots on goal instead of goals is that the

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national team over the participation of their players in the 2008 Olympic games. Similarly, Guy Demel of Hamburger SV forwent playing for his home country Ivory Coast in the Africa Cup of Nations in 2008 to have more time available for his club.

[13]The data was kindly provided by IMIPRE AG, a company specialized in collecting and selling soccer data.

[14]Their website is comunio.de.

[15]We relied on ESPNsoccernet.com, FIFA.com, Kicker.de, Worldfoot-ball.net, football-database.eu, as well as the sites of national soccer associations.

[16]Since no team ever plays twice the same day, each player-game day combination corresponds to a unique player-match combination.

Table 1: Number of players by nationality

Group	Nationality	Players
Euro 2008	Czech Republic	8
	Croatia	7
	France	2
	Germany	121
	Greece	3
	Netherlands	5
	Poland	7
	Portugal	3
	Romania	2
	Russia	1
	Sweden	2
	Turkey	3
	All Euro 2008	164
non-Euro 2008	Albania	2
	Algeria	1
	Argentinia	5
	Australia	2
	Belgium	3
	Bosnia-Herzegovina	3
	Brazil	17
	Cameroon	2
	Canada	1
	China	1
	Congo DR	1
	Denmark	7
	Egypt	1
	Finland	1
	Georgia	1
	Ghana	3
	Guinea	1
	Hungary	2
	Iran	2
	Ivory Coast	3
	Japan	1
	Macedonia	2
	Mexico	2
	Namibia	1
	Nigeria	1
	Paraguay	2
	Peru	1
	Serbia	3
	Slovakia	3
	South Africa	1
	Tunesia	2
	Uruguay	2
	USA	1
	All non-Euro 2008	81
	All players	246

Notes: The sample excludes goalkeepers, players of Austrian or Swiss nationality, or players for whom we have observations in one season only or only either after or before the official *Euro 2008* qualification date.

former occur much more frequently. It is not unusual for matches to end without any goals.

Passes received - The data contains the number of passes a player receives from his teammates in every match. This is a good indicator of how active and fit a player is, and of his teammates' trust in his ability to make a valuable contribution.

Ball contacts - Ball contacts is a more aggregate measure than passes received of how involved a player is, and also reflects a player's success in obtaining the ball.

Duels won - A duel is a situation where two players fight for the ball in direct confrontation. A duel counts as won if the player himself or one of his teammates obtains the ball in the end. Duels won measures physical fitness and dedication. Duels carry a high risk of injury, and a player who is keen on avoiding an injury may choose to fight less vigorously in a duel or stay out of duels altogether.

Minutes played - The data also include detailed information on player substitutions. Coaches are allowed to make at most three substitutions per match, and typically make use of this possibility at least twice. Approximately 80% of substitutions take place in the last 30 minutes of a match (total duration is 90 minutes plus a few minutes extra time). It makes sense to view a player's number of minutes played as a relevant output measure. First, players' performances on the field influence substitution decisions. Second, the club coach's decision to let a player be a starter or substitute him in depends on the player's effort and performance during training.[17]

The discussion of substitutions implies that observed changes in individual outputs per match could be due to changes in minutes played (see Table 3 for correlations between per match outputs and minutes played). We take two steps to disentangle other output dimensions from minutes played. First, we use outputs per minute played instead of per match to measure performance. Second, for output per minute regressions we keep only observations associated with at least 71 minutes, the median substitution time for starters conditional on being substituted out. The second restriction is useful to avoid comparing observations associated with only a few minutes on the field (usually towards the end of a match) and much longer field appearances. The average number of ball contacts per minute, for example, is about 0.606 for players who play 71 minutes or less, but 0.635 for players who play more than 71 minutes. The difference between the averages for players who play more than 71 and those who play more than 90 minutes is much smaller: 0.635 versus 0.628. Adding the condition that minutes played exceed the median substitution time for starters hence substantially alleviates the problem of comparing observations based on field appearances of different durations, while permitting us to keep

[17]Even famous players sometimes have to work hard to convince the coach to let them play. A point in case is Lukas Podolski, a star of the German national team during World Cup 2006, who had just five Bundesliga starts between August 2007 and September 2008 at the Bayern München team.

observations of players who were substituted out towards the end of a match.

In addition to the objectively measurable outputs listed so far, we use the grades that the soccer magazines *Kicker* and *Sportal* assign to players after each match as performance measures. Grades have the advantage of being an overall assessment of a player's multi-dimensional performance. The disadvantage is that grades are subjective judgements by journalists, and hence likely to be influenced by expectations prior to the match and subjective biases. Grades are recorded as numbers between 1 (excellent) and 6 (insufficient) in the data, but we used the linear transformation '6−grade' to generate a measure that is increasing and thereby facilitate the interpretation of results.

Table 2 presents summary statistics for players in the control and treatment group, respectively, and Table 3 reports correlations between the different output measures. All statistics refer to *Bundesliga* club matches.

Table 2: Summary statistics for players from nations participating (164 players) and not participating (81 players) in the *Euro 2008*

Variable	Mean	Std. Dev.	Min	Max
Euro 2008 nationalities (N = 6588)				
Age	27.08	3.91	19.43	38.58
Defense (dummy)	.339	.473	0	1
Midfield (dummy)	.485	.500	0	1
Forward (dummy)	.176	.381	0	1
Minutes played	73.87	27.25	1	96
Goals per minute	.002	.006	0	.125
Shots on goal per minute	.017	.024	0	.33
Passes received per minute	.309	.154	0	1.33
Ball contacts per minute	.613	.207	0	2
Duels won per minute	.135	.072	0	1
Kicker grade (N = 5664)	2.349	.925	0	5
Sportal grade (N = 5922)	2.421	.813	0	5
Non-*Euro 2008* nationalities (N = 3450)				
Age	28.76	3.32	19.90	36.69
Defense (dummy)	.388	.487	0	1
Midfield (dummy)	.405	.491	0	1
Forward (dummy)	.207	.405	0	1
Minutes played	74.04	26.86	1	96
Goals per minute	.002	.008	0	.25
Shots on goal per minute	.018	.026	0	.5
Passes received per minute	.313	.150	0	1.06
Ball contacts per minute	.626	.208	0	1.6
Duels won per minute	.140	.073	0	2
Kicker grade (N = 2971)	2.328	.958	0	5
Sportal grade (N = 3142)	2.445	.830	0	5

Notes: The sample excludes goalkeepers, players of Austrian or Swiss nationality, or players for whom we have observations in only one season or only either after or before the official *Euro 2008* qualification date. The summary statistics are calculated on the basis of observations associated with a positive number of minutes on the field.

Table 3: Correlations between different output measures

	Variables		Grades		Output per game					Output per minute played			
		Minutes played	Kicker	Sportal	Shots on goal	Goals	Passes received	Ball contacts	Duels won	Shots on goal	Goals	Passes received	Ball contacts
grades	Kicker	.2178	1.000										
	Sportal	.1562	.7122	1.000									
per game	Shots on goal	.2270	.2051	.2791	1.000								
	Goals	.0818	.4299	.4587	.3852	1.000							
	Passes received	.5878	.1173	.1334	.2469	.0501	1.000						
	Ball contacts	.7616	.1651	.1457	.1576	.0025	.8820	1.000					
	Duels won	.6726	.2003	.1709	.1956	.0652	.4350	.6224	1.000				
per minute played	Shots on goal	-.1413	.1757	.2414	.7079	.2788	-.0075	-.1158	-.0626	1.000			
	Goals	-.0598	.4083	.4148	.2498	.7563	-.0348	-.0827	-.0346	.3410	1.000		
	Passes received	.0148	.0335	.0652	.1287	.0045	.7292	.4883	.0603	.1107	.0129	1.000	
	Ball contacts	.1152	.0696	.0733	.0135	-.0684	.6527	.6547	.2377	.0084	-.0513	.8093	1.000
	Duels won	-.0323	.1215	.1044	.0328	.0042	.0223	.0857	.5237	.0497	.0143	.0786	.2661

Notes: The sample excludes goalkeepers, players of Austrian or Swiss nationality, or players for whom we have observations in one season only or only either after or before the official *Euro 2008* qualification date. Output per minute measures are calculated using only observations associated with a positive number of minutes on the field.

Our data also contain information about fouls. Conceptually, fouls suffered could be interpreted as a positive performance measure, the idea being that stronger players are more difficult to stop for the opponent team. Fouls committed can be viewed as a measure of destructive effort. This is the approach taken by Garicano and Palacios-Huerta (2006), who provide empirical evidence for Lazear's (1989) prediction that relative performance evaluations can lead to undesirable sabotage. Once we control for constant differences between players by means of player fixed effects, however, our regressions show no significant effects of nomination contest participation on either fouls suffered or fouls committed.

4 Empirical Strategy

To test for the effects of nomination contest participation on players with different chances of being selected for the *Euro 2008*, we first construct the following time-varying variable in $[0, 1]$ that measures player i's more recent national team history:

$$\text{pastselect}_{it} = \frac{\begin{array}{c}\text{number of } i\text{'s field appearances in the} \\ \text{past 15 matches of his nation's national team}\end{array}}{15}, \tag{2}$$

where national team matches include friendly matches, qualification matches for the *Euro 2008* or other international tournaments, and tournament matches.[18] Players' recent national team participations, as captured by pastselect, are based on national team coaches' perceptions of players' skills, which will also determine future nominations. Players with higher pastselect values should hence have greater future nomination probabilities than rival candidates with lower pastselect values. Table 12 in Appendix B shows that the values of pastselect at the time of final nomination decisions (at the end of the 07/08 season) are indeed closely related to the actual nominations for the German *Euro 2008* team. Uncertainty seems to have been greatest for players with final values of pastselect between .1 and .5: three out of ten players in this group were nominated. At high values, pastselect seems to understate a player's actual nomination chance: all players whose pastselect at the end of the 07/08 season exceeded .6 were nominated. Overall, the predicted qualitative relation between pastselect and nomination contest effort is the same

[18] The results remain similar if we treat each other tournament as consisting of a single match when constructing pastselect. The results are also robust to small changes in the number of past games used to construct pastselect, or to using the proportion of a player's appearances in either all national team matches in the past 360 days or all national team matches since summer 2005 or summer 2006 instead of the definition in (2). Only actual field appearances are used to compute pastselect because for some national team matches we were unable to obtain information on the full list of reserve players.

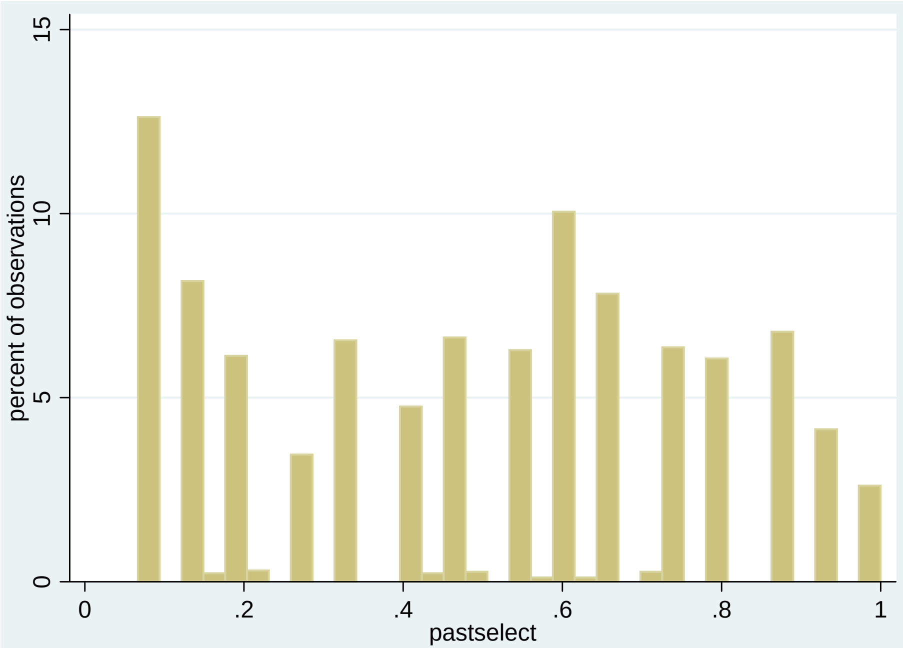

Figure 3: Histogram of pastselect$_{it}$ for nationals of *Euro 2008* nations, conditional on pastselect$_{it} > 0$. The number of players is 59, and the number of observations is 2607.

as that between nomination chance and effort, although pastselect should not be understood as a precise estimate of individual nomination probability.

In our sample, pastselect$_{it} = 0$ at all dates t for 105 out of the 164 players in the treatment group, and for 14 of the 81 players in the control group. Figure 3 depicts the distribution of pastselect observations for players of *Euro 2008* nationalities, conditional on pastselect$_{it} > 0$. Figure 4 shows the analogue to Figure 3 for the control group. The histograms confirm that the data contain variation in nomination chances. Many *Bundesliga* players are sometimes selected for their national team, but there are relatively few observations with pastselect very close to 1, which is probably due to the fact that most soccer superstars work for better-paying English, Spanish or Italian clubs.

Our theory predicts that nomination contest participation affects the effort decision of players who have a positive nomination chance. Players in the treatment group who currently believe they will be nominated with an intermediate probability should have the strongest incentives to exert additional effort in order to impress the national coach. A player whose current nomination chance is close to one, on the other hand, expecting that a small performance change will not affect the national coach's decision, has weaker incentives to exert additional effort. In addition, players with positive nomination chances should have stronger than normal incentives to avoid exhaustion and injuries prior to the Euro Cup, which could even lead to a

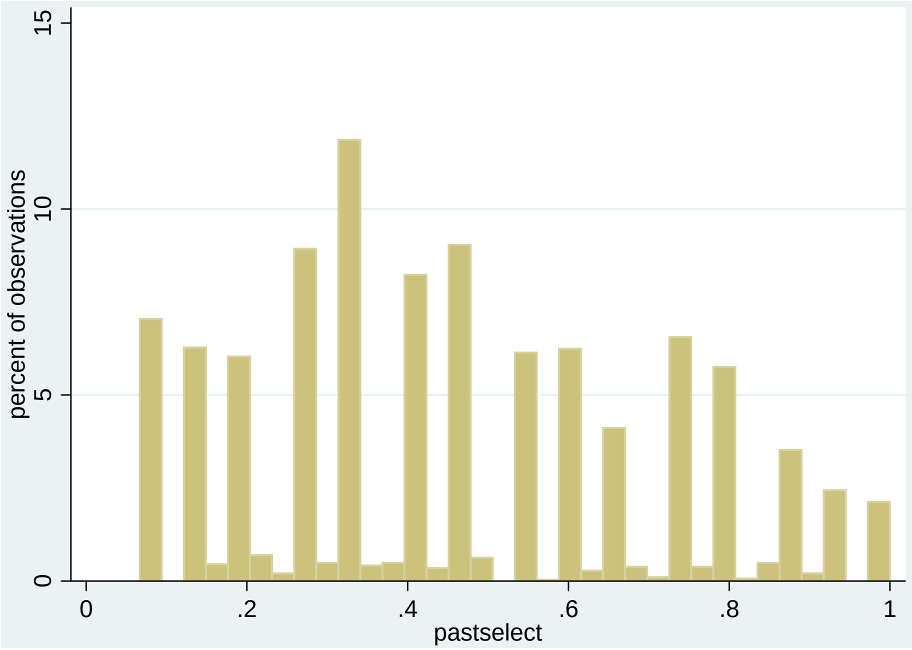

Figure 4: Histogram of pastselect_{it} for players who are not from *Euro 2008* nations, conditional on $\text{pastselect}_{it} > 0$. The number of players is 67, and the number of observations is 2867.

negative net effect of contest participation for players with high nomination chances. To test these predictions, we run the following difference-in-difference-in-differences regressions:

$$\begin{aligned} Y_{int} = \quad & \delta_0 \text{qualified}_{nt} \\ & + \delta_1 \text{qualified}_{nt} \text{pastselect}_{it} + \delta_2 \text{qualified}_{nt} \text{pastselect}_{it} \left(1 - \text{pastselect}_{it}\right) \\ & + \eta_1 \text{pastselect}_{it} + \eta_2 \text{pastselect}_{it} \left(1 - \text{pastselect}_{it}\right) \\ & + \rho_1 \text{euro}_n \text{pastselect}_{it} + \rho_2 \text{euro}_n \text{pastselect}_{it} \left(1 - \text{pastselect}_{it}\right) \\ & + \pi_1 \text{post}_t \text{pastselect}_{it} + \pi_2 \text{post}_t \text{pastselect}_{it} \left(1 - \text{pastselect}_{it}\right) \\ & + \gamma_i + \alpha_t + X_{it}' \beta + \varepsilon_{int}. \end{aligned} \tag{3}$$

where Y_{int} is the output of player i of nationality n on game day t. We run separate regressions for different output measures. The treatment dummy qualified_{nt} equals 1 if and only if nation n is qualified for the *Euro 2008* at time t. The theory predicts that δ_0, the treatment effect for players with no recent national team participations, is zero. The coefficient δ_2 is predicted to be positive, since players with uncertain chances, i.e., high values of pastselect(1-pastselect), have strong effort incentives. δ_1 is predicted to be negative if injury concerns and energy preservation strategies are relevant. We also run regressions with tertile or quartile dummies of pastselect instead of pastselect and pastselect(1-pastselect) as robustness checks for the functional form assumption. In all cases, the various pastselect variables also enter the regression equations interacted with a euro_n dummy that indicates whether nation n was a *Euro 2008* participant, and post_t, which indicates the time period after the official *Euro 2008* qualification date (November 21, 2007).[19]

The player fixed effects γ_i pick up (time-invariant) skill differences between players, and the game day fixed effects α_t control for changes in playing conditions over time that affect all clubs. X_{it} also includes dummies that indicate the club the player currently works for,[20] and dummies that indicate the opponent team i's club faces on day t. Since it is relatively common for players to occupy different field positions (forward, midfield or defense) in different matches, the covariates moreover include field position dummies. Finally, X_{it} includes a homegame_{it} dummy indicating whether i's current club plays in its home stadium on day t, and an unfit_{it} dummy indicating whether the player is injured or recovering from an injury.[21]

In our main alternative specification, we use club-game day dummies instead of the game day, club, opponent, and homegame dummies. There are two club-game day dummies per match, one per participating club. These dummies capture unobserved differences in the marginal returns from a victory across matches and clubs (depending, for example, on the current degree of competition for the championship and the club's current ranking), and other differences in playing conditions (weather etc.) between matches and clubs. Inclusion of these finer club-game day dummies substantially improves fit.

The identifying assumption is that in the absence of the Euro Cup treatment, players from qualified and from non-qualified nations would have evolved similarly over time (given controls). Since players in the treatment and the control group work in the same environment and are subject to similar incentive systems in the absence of international Cups, we find little reason to doubt this. A player's eligibility for the Euro Cup treatment, i.e., his nation's participation in the Euro Cup qualifications, is determined exogenously by geography and the player's nationality.[22] Within the group of Europeans, the assignment of the treatment, i.e., a nation's qualification, should depend on the skills of the players who participated in the Euro Cup qualification matches, so for a small number of European players selection into the treatment group is not completely random at this stage. Since we control for constant output differences by means of player fixed effects, however, bias caused by potential correlation between these players' outputs and treatment

[19] $\text{qualified}_{nt} = \text{post}_t \times \text{euro}_n$ for nations that qualified on the official qualification date (November 21, 2007). For nations that already de facto qualified at an earlier date, qualified_{nt} is equal to 1 from the de facto qualification date onwards.

[20] Several players in our sample switched between clubs in the sample period.

[21] Note that if a player is seriously injured, he will not show up in our output dataset, which only contains observations for players who were either on the reserve bench or on the field.

[22] In rare cases players change nationality. Formerly Brazilian player Deco's adopted Portuguese citizenship, for example, mainly to participate in the Euro 2004 and World Cup 2006. Authorities and the FIFA have a critical attitude to such steps, however, which are therefore very rare.

status is largely if not completely eliminated in our results.

An underlying assumption is that the de facto qualification dates are relevant for determining the beginning of the treatment for *Euro 2008* - Europeans. Our analysis builds on the insight that on a nation's de facto qualification date its qualification probability exhibits a discrete and permanent upward jump (to one).[23] One may argue however that players from countries that are likely to qualify may have already altered their effort earlier on. Such effects tend to bias against finding performance responses to qualification, thereby making our estimates conservative.

Because the data on minutes played take on nonnegative integer values (between 0 and 96), a count model is appropriate in regressions with minutes played as the dependent variable. We will use the negative binomial model, as the Poisson model is rejected at high degrees of confidence.[24] For the other dependent variables, outputs (shots on goal,...) per minute played and grades, we use OLS estimation. Standard errors are robust and clustered at the individual player level to take into account serial correlation.[25] The resulting estimator of the variance-covariance matrix is consistent as the number of players in our data is large (see Bertrand et al., 2004).

5 Results

Tables 4 to 11 report results of regressions with different performance measures as the dependent variable. For minutes played, we present both OLS and negative binomial regression results (Tables 4 and 5). We will first discuss overall patterns in the results, and then turn to differences between various output measures.

Columns (1) and (2) of each table report results for the basic regression specification in equation (3). The regressors of main interest are the interactions qualified$\times$pastselect(1-pastselect) and qualified$\times$pastselect. For all output measures, the coefficient of the former is positive and that of the latter negative, as predicted by our theory. For minutes played, passes received, ball contacts, and *Sportal* grades, both coefficients are statistically significant, mostly at the 1% or 5%

[23]Similarly, for non-qualified European nations there is a downward jump to zero at some point in time, in some cases long before the official qualification date. The group of players from such nations in our sample is small ($n = 23$).

[24]Allison and Waterman (2002) and Guimarães (2008) show that for the negative binomial model the estimator proposed by Hausman et al. (1984) is a conditional fixed effects estimator under very specific assumptions only. As suggested by Allison and Waterman (2002), player fixed effects can be included by means of player dummies, however, which is the approach we follow.

[25]If class is player identity, the intraclass correlations for the various output measures we employ lie between 0.2 and 0.4. Note also that while the regression equation in the text allow error terms to depend on nationality n, within-group correlations at the nationality level are low: for all our output measures the intraclass correlation if class is nationality lies below 0.1, in many cases even below 0.05.

level. The sizes of the coefficients are such that the implied net impact of nomination contest participation is positive for pastselect values up to somewhere between .6 and .7, depending on the output measure, and negative thereafter. The latter is in line with our earlier observation, based on Table 12 in Appendix B, that pastselect above .6 suggests certain nomination, so that injury concerns dominate. Positive effects are maximal for pastselect between .3 and .4, i.e., for players with appearances in $30-40\%$ of their country's recent national team matches. For instance, the estimated effect of nomination contest participation on the passes received per minute of a player with pastselect equal to .3 is about +8% (with respect to pre-treatment observations with pastselect-values between .2 and .4 of treatment group players). The corresponding effects on other performance measures are of similar magnitudes: +7% for ball contacts per minute, +9% (or 5.7 field minutes) for minutes played, and +6% for Sportal grades. For shots on goal (Table 8), where only the positive interaction qualified×pastselect(1-pastselect) is significant ($p < 0.1$), the estimated positive impact of nomination contest participation for a player with pastselect $= 0.3$ is as high as 25%.[26] The coefficient of qualified, which measures the impact of nomination contest participation for players without any recent national team participations, is insignificant in all these regressions. This is consistent with the theoretical prediction that players without nomination chances do not alter their efforts.

Columns (3) to (6) of the regression tables report results of regression with dummies for different percentiles of positive pastselect values. These regressions confirm that negative effects for players with high nomination chances are not an artifact of the functional form of pastselect in the basic regression equation discussed so far. In the regressions with club-game day dummies, interactions of the treatment with the top tertile or quartile of pastselect (pastselect above .6429 and .7333, respectively) have a significant negative impact on many output measures: minutes played, ball contacts, passes received, *Kicker* grades, and duels won. These negative effects are economically significant. The regressions for ball contacts per minute with club-game day dummies (columns (3) and (5) in Table 7) imply output reductions of about 10% for players in the top tertile and top quartile. The corresponding effects on passes received per minute are −14% and −13%.

For low pastselect percentiles, the coefficients of the interactions with the treatment are generally positive,[27] as predicted by the theory, but not always significant.

[26] All effects were calculated on the basis of the regressions with club-game day dummies in columns (1). Since the regressions results with different dummies reported in columns (2) are very similar, the estimated effects would be very close if we used those estimated instead.

[27] An exception occurs in Table 11 where in column (5) the interaction of qualified with the lowest pastselect quartile is negative and significant. The coefficient of qualified is positive and significant in this regression as well, however, and jointly the two coefficients are statistically insignificant. For observations in the highest pastselect quartile, on the other hand, the joint effect is is negative and significant at the 5% level.

Where significant, the effects are substantial. In the case of passes received (Table 6), for instance, we find positive effects of about 11% and 9% for the lowest pastselect tertile and the second pastselect quartile.

In the regressions with club, opponent and game day dummies (columns (4) and (6) of each table), the coefficient of qualified is negative and statistically significant for some output measures, which is inconsistent with the theoretical prediction that nomination contest participation affects only the effort of players with positive nomination chances. In all regressions, however, the effect vanishes once finer club-game day dummies are used.

There are interesting differences between the findings for the various output measures. The theory implies that players with high nomination chances should reduce activities that carry a high injury risk. This is consistent with our finding that nomination contest participation has negative effects on the number of duels won. In the basic regression equation (columns (1) and (2) in Table 9) only the negative interaction term with pastselect is significant and in the regressions with dummies the only significant effects are negative ones. These negative impacts are economically significant: we find a 13% reduction in the number of duels won for players in the top pastselect tertile for example. Players with high nomination chances hence seem to be less persistent in duels, which carry a much higher injury risk than actions in less direct confrontation with players of the opponent team. The results are similar in unreported regressions with total duels instead of duels won as the dependent variable, which suggests that players with high nomination chances also avoid fighting duels in the first place.

The control variables have the expected signs. The coefficient of homegame is positive and highly significant in most regressions. Interestingly, homegame is also significant in the regressions with grades as the dependent variable. Soccer journalists hence do not seem to discount performances for the well-known homegame advantage when grading players. A forward field position is associated with more frequent goal attempts but fewer ball contacts, while midfield positions are associated with significantly more duels than forward or defense positions. The results for minutes played show that there are also more substitutions of players in forward and midfield positions than of players in defense positions. The coefficient of injured has a negative sign in all regressions, but is statistically significant for minutes played only.

To summarize our findings on the differential effects of the Euro Cup treatment:

1. Players from qualified countries with intermediate national team nomination chances perform better in club matches (relative to players of other nationalities with similar national team experience) after their nations' qualifications for the *Euro 2008* than before.

2. Players from qualified countries with very high national team nomination

Table 4: Regression results for minutes played (Negbin FE Model)

VARIABLES	Minutes played					
	(1)	(2)	(3)	(4)	(5)	(6)
qualified	.048 (.102)	.067 (.085)	.059 (.105)	.078 (.088)	.062 (.106)	.079 (.088)
qualified ×						
pastselect	-.632*** (.241)	-.538*** (.196)				
pastselect(1-pastselect)	1.659** (.814)	1.426** (.675)				
qualified ×						
$\text{pastselect}_{1stTertile}$			.076 (.165)	.059 (.146)		
$\text{pastselect}_{2ndTertile}$			.152 (.163)	.133 (.111)		
$\text{pastselect}_{3rdTertile}$			-.273** (.134)	-.264** (.106)		
qualified ×						
$\text{pastselect}_{1stQuartile}$					.049 (.188)	-.00002 (.160)
$\text{pastselect}_{2ndQuartile}$					.269 (.170)	.254* (.134)
$\text{pastselect}_{3rdQuartile}$					-.140 (.150)	-.039 (.132)
$\text{pastselect}_{4thQuartile}$					-.336** (.170)	-.269** (.105)
forward	-.576*** (.144)	-.508*** (.128)	-.579*** (.143)	-.510*** (.128)	-.580*** (.143)	-.507*** (.128)
midfield	-.372*** (.100)	-.326*** (.089)	-.372*** (.101)	-.325*** (.089)	-.371*** (.101)	-.321*** (.090)
injured	-.181*** (.044)	-.155*** (.035)	-.178*** (.044)	-.154*** (.035)	-.177*** (.044)	-.153*** (.035)
pastselect	.144 (.198)	.245 (.151)				
pastselect(1-pastselect)	-.035 (.635)	.252 (.564)				
Gameday-club FE	Yes		Yes		Yes	
Gameday FE		Yes		Yes		Yes
Club FE		Yes		Yes		Yes
Opponent FE		Yes		Yes		Yes
Observations	11799	11799	11799	11799	11799	11799

Notes: The table reports negative binomial regression estimates. Values between parentheses are robust standard errors clustered at the player level. Only observations from players who are neither goalkeepers nor Austrian or Swiss are included. Moreover, the sample includes only players who were active in both the 06/07 and the 07/08 season, and before and after 21 Nov 2007, and with at least one strictly positive observation of minutes played in the two seasons.

*** p< 0.01, ** p< 0.05, * p < 0.1

Table 5: Regression results for minutes played (Linear FE Model)

VARIABLES	Minutes played					
	(1)	(2)	(3)	(4)	(5)	(6)
qualified	2.663 (3.587)	4.280 (3.172)	3.532 (3.624)	5.238 (3.213)	3.626 (3.628)	5.337* (3.219)
qualified ×						
pastselect	-33.490*** (11.240)	-29.670*** (10.660)				
pastselect(1-pastselect)	75.180** (35.950)	62.680* (34.200)				
qualified ×						
$\text{pastselect}_{1stTertile}$			1.688 (6.309)	-.324 (6.116)		
$\text{pastselect}_{2ndTertile}$			4.489 (6.503)	4.002 (4.980)		
$\text{pastselect}_{3rdTertile}$			-16.850*** (5.426)	-17.200*** (5.223)		
qualified ×						
$\text{pastselect}_{1stQuartile}$					.279 (7.103)	-2.777 (6.631)
$\text{pastselect}_{2ndQuartile}$					9.538 (7.295)	9.267 (6.734)
$\text{pastselect}_{3rdQuartile}$					-3.888 (6.940)	-4.310 (6.552)
$\text{pastselect}_{4thQuartile}$					-19.480** (5.695)	-18.250*** (4.926)
forward	-21.960*** (4.742)	-20.810*** (4.622)	-22.120*** (4.733)	-20.910*** (4.623)	-22.120*** (4.732)	-20.800*** (4.620)
midfield	-17.480*** (3.482)	-16.000*** (3.406)	-17.550*** (3.491)	-16.000*** (3.423)	-17.430*** (3.513)	-15.850*** (3.434)
injured	-8.572*** (1.686)	-7.636*** (1.488)	-8.610*** (1.687)	-7.693*** (1.470)	-8.530*** (1.695)	-7.616*** (1.486)
pastselect	6.153 (9.091)	11.670 (7.603)				
pastselect(1-pastselect)	-9.026 (26.740)	-1.078 (25.910)				
Gameday-club FE	Yes		Yes		Yes	
Gameday FE		Yes		Yes		Yes
Club FE		Yes		Yes		Yes
Opponent FE		Yes		Yes		Yes
Observations	11799	11799	11799	11799	11799	11799
Variance captured by player FE	.58	.56	.58	.56	.58	.56
R^2	.13	.07	.13	.07	.13	.07

Notes: The table reports linear fixed effects regression estimates. Values between parentheses are robust standard errors clustered at the player level. Only observations from players who are neither goalkeepers nor Austrian or Swiss are included. Moreover, the sample includes only players who were active in both the 06/07 and the 07/08 season, and before and after 21 Nov 2007, and with at least one strictly positive observation of minutes played in the two seasons.
*** p< 0.01, ** p< 0.05, * p < 0.1

Table 6: Regression results for passes received

VARIABLES	Passes received per minute played					
	(1)	(2)	(3)	(4)	(5)	(6)
qualified	-.011 (.009)	-.020 (.012)	-.012 (.009)	-.027** (.011)	-.011 (.009)	-.026** (.011)
qualified ×						
pastselect	-.131*** (.034)	-.099** (.040)				
pastselect(1-pastselect)	.322*** (.099)	.356*** (.115)				
qualified ×						
$\text{pastselect}_{1stTertile}$			.036** (.017)	.066*** (.017)		
$\text{pastselect}_{2ndTertile}$			-.006 (.018)	.015 (.016)		
$\text{pastselect}_{3rdTertile}$			-.053*** (.022)	-.014 (.020)		
qualified ×						
$\text{pastselect}_{1stQuartile}$					.026 (.017)	.056*** (.018)
$\text{pastselect}_{2ndQuartile}$					.030* (.017)	.054*** (.018)
$\text{pastselect}_{3rdQuartile}$					-.034* (.018)	.001 (.020)
$\text{pastselect}_{4thQuartile}$					-.049** (.023)	-.010 (.027)
home game		.030*** (.003)		.030*** (.003)		.030*** (.003)
forward	-.005 (.012)	-.009 (.013)	-.005 (.012)	-.010 (.013)	-.006 (.012)	-.008 (.013)
midfield	.009 (.009)	.011 (.010)	.009 (.009)	.010 (.010)	.009 (.009)	.010 (.010)
injured	-.003 (.006)	-.001 (.007)	-.003 (.006)	-.001 (.006)	-.003 (.006)	-.001 (.007)
pastselect	-.009 (.018)	.006 (.025)				
pastselect(1-pastselect)	.017 (.057)	.065 (.072)				
Gameday-club FE	Yes		Yes		Yes	
Gameday FE		Yes		Yes		Yes
Club FE		Yes		Yes		Yes
Opponent FE		Yes		Yes		Yes
Observations	6747	6747	6747	6747	6747	6747
Variance captured by player FE	.46	.38	.50	.38	.48	.38
R^2	.55	.19	.55	.19	.55	.19

Notes: The table reports linear fixed effects estimates. Values between parentheses are robust standard errors clustered at the player level. Only observations associated with more than 71 minutes played and of players who are neither goalkeepers nor Austrian or Swiss are included. Moreover, the sample includes only players who were active in both the 06/07 and the 07/08 season (before and after 21 Nov 2007), and with at least one strictly positive observation of the dependent variable in these two seasons.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Table 7: Regression results for ball contacts

VARIABLES	Ball contacts per minute played					
	(1)	(2)	(3)	(4)	(5)	(6)
qualified	-.016 (.012)	-.024 (.016)	-.018 (.012)	-.033** (.015)	-.018 (.012)	-.032** (.015)
qualified ×						
pastselect	-.183*** (.045)	-.117** (.047)				
pastselect(1-pastselect)	.455*** (.126)	.410*** (.136)				
qualified ×						
$\text{pastselect}_{1stTertile}$			.051** (.020)	.079*** (.021)		
$\text{pastselect}_{2ndTertile}$			-.009 (.023)	.010 (.023)		
$\text{pastselect}_{3rdTertile}$			-.075** (.031)	-.016 (.030)		
qualified ×						
$\text{pastselect}_{1stQuartile}$					.038* (.022)	.070*** (.021)
$\text{pastselect}_{2ndQuartile}$					.035 (.022)	.052** (.024)
$\text{pastselect}_{3rdQuartile}$					-.048 (.028)	-.003 (.028)
$\text{pastselect}_{4thQuartile}$					-.070** (.032)	-.010 (.032)
home game		.037*** (.004)		.037*** (.004)		.037*** (.004)
forward	-.112*** (.020)	-.118*** (.020)	-.112*** (.020)	-.120*** (.021)	-.102*** (.034)	-.118*** (.020)
midfield	-.083*** (.017)	-.081*** (.017)	-.083*** (.018)	-.082*** (.017)	-.057** (.027)	-.082*** (.018)
injured	-.008 (.007)	-.001 (.008)	-.007 (.007)	-.001 (.008)	-.147 (.013)	-.001 (.008)
pastselect	-.015 (.026)	.009 (.031)				
pastselect(1-pastselect)	.041 (.078)	.092 (.085)				
Gameday-club FE	Yes		Yes		Yes	
Gameday FE		Yes		Yes		Yes
Club FE		Yes		Yes		Yes
Opponent FE		Yes		Yes		Yes
Observations	6747	6747	6747	6747	6747	6747
Variance captured by player FE	.54	.47	.56	.47	.54	.47
R^2	.46	.17	.46	.17	.46	.17

Notes: The table reports linear fixed effects estimates. Values between parentheses are robust standard errors clustered at the player level. Only observations associated with more than 71 minutes played and of players who are neither goalkeepers nor Austrian or Swiss are included. Moreover, the sample includes only players who were active in both the 06/07 and the 07/08 season (before and after 21 Nov 2007), and with at least one strictly positive observation of the dependent variable in these two seasons.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Table 8: Regression results for shots on goal

VARIABLES	Shots on goal per minute played					
	(1)	(2)	(3)	(4)	(5)	(6)
qualified	-.001 (.001)	-.001 (.001)	-.001 (.001)	-.001 (.001)	-.001 (.001)	-.001 (.001)
qualified ×						
pastselect	-.005 (.004)	-.005 (.003)				
pastselect(1-pastselect)	.023* (.013)	.022* (.012)				
qualified ×						
$\text{pastselect}_{1stTertile}$			.001 (.002)	.001 (.002)		
$\text{pastselect}_{2ndTertile}$			.001 (.003)	.001 (.002)		
$\text{pastselect}_{3rdTertile}$			.0002 (.003)	.001 (.002)		
qualified ×						
$\text{pastselect}_{1stQuartile}$					.001 (.003)	.0001 (.002)
$\text{pastselect}_{2ndQuartile}$					.004 (.003)	.004 (.003)
$\text{pastselect}_{3rdQuartile}$					.002 (.003)	.002 (.003)
$\text{pastselect}_{4thQuartile}$					-.005 (.003)	-.004 (.003)
home game		.036*** (.0003)		.004*** (.0003)		.004*** (.0003)
forward	.008*** (.002)	.008*** (.002)	.008*** (.002)	.008*** (.002)	.008*** (.002)	.008*** (.002)
midfield	.007*** (.001)	.007*** (.001)	.007*** (.001)	.007*** (.001)	.007** (.001)	.007*** (.001)
injured	-.002* (.001)	-.002 (.001)	-.002 (.001)	-.002 (.001)	-.002 (.001)	-.001 (.001)
pastselect	.005* (.003)	.005** (.002)				
pastselect(1-pastselect)	.0004 (.008)	.001 (.008)				
Gameday-club FE	Yes		Yes		Yes	
Gameday FE		Yes		Yes		Yes
Club FE		Yes		Yes		Yes
Opponent FE		Yes		Yes		Yes
Observations	6747	6747	6747	6747	6747	6747
Variance captured by player FE	.41	.41	.42	.40	.41	.41
R^2	.24	.07	.24	.07	.24	.07

Notes: The table reports linear fixed effects estimates. Values between parentheses are robust standard errors clustered at the player level. Only observations associated with more than 71 minutes played and of players who are neither goalkeepers nor Austrian or Swiss are included. Moreover, the sample includes only players who were active in both the 06/07 and the 07/08 season (before and after 21 Nov 2007), and with at least one strictly positive observation of the dependent variable in these two seasons.

*** p < 0.01, ** p < 0.05, * p < 0.1

Table 9: Regression results for duels won

VARIABLES	Duels won per minute played					
	(1)	(2)	(3)	(4)	(5)	(6)
qualified	.0001 (.004)	.002 (.004)	.0001 (.004)	.001 (.004)	.0002 (.004)	.002 (.004)
qualified ×						
pastselect	-.024** (.010)	-.022** (.009)				
pastselect(1-pastselect)	.032 (.039)	.023 (.036)				
qualified ×						
$\text{pastselect}_{1stTertile}$			.003 (.008)	.005 (.007)		
$\text{pastselect}_{2ndTertile}$			-.009 (.007)	-.010 (.006)		
$\text{pastselect}_{3rdTertile}$			-.017** (.007)	-.014*** (.006)		
qualified ×						
$\text{pastselect}_{1stQuartile}$					.002 (.008)	.004 (.007)
$\text{pastselect}_{2ndQuartile}$					-.008 (.008)	-.008 (.007)
$\text{pastselect}_{3rdQuartile}$					-.008 (.008)	-.010 (.007)
$\text{pastselect}_{4thQuartile}$					-.014* (.008)	-.012* (.007)
home game		.005*** (.001)		.005*** (.001)		.005*** (.001)
forward	.008 (.006)	.007 (.005)	.008 (.006)	.007 (.005)	.009 (.006)	.008 (.005)
midfield	.009** (.004)	.009** (.004)	.008** (.004)	.009** (.004)	.009** (.004)	.009** (.004)
injured	-.002 (.003)	-.002 (.003)	-.002 (.003)	-.002 (.003)	-.002 (.003)	-.002 (.003)
pastselect	-.003 (.009)	-.006 (.007)				
pastselect(1-pastselect)	-.011 (.023)	-.001 (.023)				
Gameday-club FE	Yes		Yes		Yes	
Gameday FE		Yes		Yes		Yes
Club FE		Yes		Yes		Yes
Opponent FE		Yes		Yes		Yes
Observations	6747	6747	6747	6747	6747	6747
Variance captured by player FE	.41	.38	.41	.38	.46	.38
R^2	.24	.07	.25	.07	.25	.07

Notes: The table reports linear fixed effects estimates. Values between parentheses are robust standard errors clustered at the player level. Only observations associated with more than 71 minutes played and of players who are neither goalkeepers nor Austrian or Swiss are included. Moreover, the sample includes only players who were active in both the 06/07 and the 07/08 season (before and after 21 Nov 2007), and with at least one strictly positive observation of the dependent variable in these two seasons.
*** p < 0.01, ** p < 0.05, * p < 0.1

Table 10: Regression results for Sportal grades

VARIABLES	Sportal grade					
	(1)	(2)	(3)	(4)	(5)	(6)
qualified	-.032 (.056)	-.071 (.058)	-.035 (.056)	-.093 (.062)	-.038 (.056)	-.092 (.062)
qualified ×						
pastselect	-.431** (.189)	-.440** (.196)				
pastselect(1-pastselect)	1.295* (.658)	1.599** (.646)				
qualified ×						
$\text{pastselect}_{1stTertile}$			.070 (.106)	.189* (.112)		
$\text{pastselect}_{2ndTertile}$			.072 (.110)	.074 (.118)		
$\text{pastselect}_{3rdTertile}$			-.140 (.123)	-.034 (.136)		
qualified ×						
$\text{pastselect}_{1stQuartile}$					.044 (.115)	.135 (.134)
$\text{pastselect}_{2ndQuartile}$					.020 (.140)	.224* (.131)
$\text{pastselect}_{3rdQuartile}$					.062 (.120)	-.004 (.138)
$\text{pastselect}_{4thQuartile}$					-.162 (.147)	-.102 (.163)
home game		.222*** (.017)		.222*** (.017)		.222*** (.017)
forward	.115 (.011)	.095 (.116)	.120 (.107)	.092 (.114)	.123 (.107)	.099 (.116)
midfield	.081 (.060)	.102 (.066)	.087 (.059)	.102 (.065)	.084 (.059)	.101 (.065)
injured	-.030 (.042)	-.054 (.044)	-.031 (.042)	-.057 (.044)	-.028 (.043)	-.053 (.044)
pastselect	-.055 (.119)	-.068 (.133)				
pastselect(1-pastselect)	.021 (.448)	-.280 (.469)				
Gameday-club FE	Yes		Yes		Yes	
Gameday FE		Yes		Yes		Yes
Club FE		Yes		Yes		Yes
Opponent FE		Yes		Yes		Yes
Observations	6721	6721	6721	6721	6721	6721
Variance captured by player FE	.32	.26	.32	.27	.33	.27
R^2	.45	.05	.45	.06	.45	.06

Notes: The table reports linear fixed effects estimates. Values between parentheses are robust standard errors clustered at the player level. Only observations associated with more than 71 minutes played and of players who are neither goalkeepers nor Austrian or Swiss are included. Moreover, the sample includes only players who were active in both the 06/07 and the 07/08 season (before and after 21 Nov 2007), and with at least one strictly positive observation of the dependent variable in these two seasons.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Table 11: Regression results for Kicker grades

VARIABLES	Kicker grade					
	(1)	(2)	(3)	(4)	(5)	(6)
qualified	-.107	-.407	.129	.032	.132*	.030
	(.075)	(.075)	(.078)	(.079)	(.076)	(.079)
qualified ×						
pastselect	-.447**	-.437*				
	(.192)	(.224)				
pastselect(1-pastselect)	.417	1.050				
	(.700)	(.767)				
qualified ×						
$\text{pastselect}_{1stTertile}$			-.172	.015		
			(.118)	(.130)		
$\text{pastselect}_{2ndTertile}$			-.105	.003		
			(.138)	(.143)		
$\text{pastselect}_{3rdTertile}$			-.335**	-.159		
			(.137)	(.155)		
qualified ×						
$\text{pastselect}_{1stQuartile}$					-.230*	-.054
					(.139)	(.158)
$\text{pastselect}_{2ndQuartile}$					-.097	.179
					(.141)	(.157)
$\text{pastselect}_{3rdQuartile}$					-.162	-.093
					(.140)	(.156)
$\text{pastselect}_{4thQuartile}$					-.355**	-.223
					(.150)	(.176)
home game		.233***		.233***		.233***
		(.020)		(.020)		(.020)
forward	.032	.045	.038	.046	.043	.048
	(.095)	(.115)	(.094)	(.113)	(.093)	(.114)
midfield	-.080	-.014	-.079	-.015	-.081	-.017
	(.054)	(.068)	(.053)	(.066)	(.053)	(.067)
injured	.017	-.018	.015	-.021	.013	-.021
	(.051)	(.055)	(.050)	(.055)	(.050)	(.055)
pastselect	.084	.040				
	(.141)	(.152)				
pastselect(1-pastselect)	-.266	-.251				
	(.450)	(.568)				
Gameday-club FE	Yes		Yes		Yes	
Gameday FE		Yes		Yes		Yes
Club FE		Yes		Yes		Yes
Opponent FE		Yes		Yes		Yes
Observations	6722	6722	6722	6722	6722	6722
Variance captured by player FE	.31	.28	.31	.27	.30	.28
R^2	.48	.05	.48	.06	.48	.06

Notes: The table reports linear fixed effects estimates. Values between parentheses are robust standard errors clustered at the player level. Only observations associated with more than 71 minutes played and of players who are neither goalkeepers nor Austrian or Swiss are included. Moreover, the sample includes only players who were active in both the 06/07 and the 07/08 season (before and after 21 Nov 2007), and with at least one strictly positive observation of the dependent variable in these two seasons.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

chances perform worse in club matches (relative to players of other nationalities with similar national team experience) after their nations' qualifications for the *Euro 2008* than before.

6 Conclusion

Contest-style rivalry, whether based on pre-specified performance criteria or reputations as the nominations for national soccer teams, arises in many contexts. Some firms explicitly offer promotion prospects or use relative performance evaluation schemes in order to provide incentives to employees. In many other situations, the principal's goal is to select the most skillful agent, but this creates similar incentives. In either case, economic theory predicts that agents' effort responses should depend on their anticipated winning probabilities. In particular, agents with intermediate winning probabilities should exert higher than normal effort. This paper provides empirical evidence for this prediction. We show that players from nations qualified for the *Euro 2008* who had been called upon by the national coach in some but not too many past national team matches improved their club performance, relative to players of other nationalities with a similar standing in their national teams, after their countries' qualifications. For players without any past national team nominations, on the other hand, there is no evidence of any improvement relative to players of other nationalities.

Moreover, we find that players who were already quite certain of their Euro Cup participations performed worse along several dimensions than they would have in the absence of the upcoming Cup. Our explanation is that these players wanted to avoid injuries and more generally preserve their strength and fitness for the Cup. Hence, while clubs often benefit from the national team nomination contests, they may actually suffer losses in the case of top players. Similar effects can occur in other situations where agents compete for a position that requires future effort instead of a monetary prize. Consider promotion contests in firms for example. An employee who expects an almost certain promotion into a different unit may be inclined to exert less effort in his current position in order to preserve energy for his new position. Such behavior inflicts a loss on the employee's current unit. Ensuring that rivalry between candidates persists is key to avoiding such losses and promoting effort. Effort will be higher if several candidates perceive that they have realistic but less than perfect chances of obtaining the promotion.

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Appendix A: Analysis of the Theoretical Model

Denote by (a_1^*, a_2^*) the equilibrium effort levels. Thanks to our normality and independence assumptions, the learning process about each agent's skill is well-known. Given the principal anticipates effort level a_j^*, the posterior distribution of η_j after observing y_j will be normal with mean

$$\frac{h_j m_j + h_\varepsilon (y_j - a_j^*)}{h_j + h_\varepsilon} \tag{A.1}$$

and precision $h_j + h_\varepsilon$.

Let us now consider j's effort decision at the beginning of the period. From (A.1) it follows that, given $a_k = a_k^*$, if j chooses a_j then he will have a higher posterior reputation than agent k with probability

$$\Pr\left\{\frac{h_j m_j + h_\varepsilon(\eta_j + a_j + \varepsilon_j - a_j^*)}{h_j + h_\varepsilon} > \frac{h_k m_k + h_\varepsilon(\eta_k + \varepsilon_k)}{h_k + h_\varepsilon}\right\} \tag{A.2}$$

$$= \Pr\left\{\frac{h_\varepsilon}{h_j + h_\varepsilon}(a_j - a_j^*) > \frac{h_k m_k + h_\varepsilon(\eta_k + \varepsilon_k)}{h_k + h_\varepsilon} - \frac{h_j m_j + h_\varepsilon(\eta_j + \varepsilon_j)}{h_j + h_\varepsilon}\right\}. \tag{A.3}$$

Define the random variable

$$\zeta_j \equiv \frac{h_k m_k + h_\varepsilon(\eta_k + \varepsilon_k)}{h_k + h_\varepsilon} - \frac{h_j m_j + h_\varepsilon(\eta_j + \varepsilon_j)}{h_j + h_\varepsilon}.$$

Our independence and normality assumptions imply that the *prior* distribution of ζ_j is normal with mean

$$z_j \equiv m_k - m_j \tag{A.4}$$

and variance[28]

$$\sigma^2 \equiv \left(\frac{h_\varepsilon}{h_k + h_\varepsilon}\right)^2 \left(\frac{1}{h_k} + \frac{1}{h_\varepsilon}\right) + \left(\frac{h_\varepsilon}{h_j + h_\varepsilon}\right)^2 \left(\frac{1}{h_j} + \frac{1}{h_\varepsilon}\right) \tag{A.5}$$

We denote this distribution by $\varphi_j(\cdot)$ with c.d.f. $\phi_j(\cdot)$. Moreover, let us denote by $\sigma(h_j, h_k, h_\varepsilon)$, equal to the square root of σ^2 defined in (A.5), the standard deviation of the distributions $\varphi_1(\cdot)$ and $\varphi_2(\cdot)$.

[28] Since the prior distributions of ζ_1 and ζ_2 have the same variance, we can simply denote this variance by σ^2, not using any subscript.

Using the newly defined variable ζ_j, the probability in (A.3) that j's posterior reputation exceeds that of k can be rewritten as

$$\Pr\left\{\zeta_j < \frac{h_\varepsilon}{h_j + h_\varepsilon}(a_j - a_j^*)\right\} = \phi_j\left(\frac{h_\varepsilon}{h_j + h_\varepsilon}(a_j - a_j^*)\right). \tag{A.6}$$

Given $a_k = a_k^*$, the marginal impact of a_j on the probability that j has a higher posterior reputation than k is equal to the first derivative of (A.6) with respect to a_j:

$$\varphi_j\left(\frac{h_\varepsilon}{h_j + h_\varepsilon}(a_j - a_j^*)\right)\frac{h_\varepsilon}{h_j + h_\varepsilon}.$$

The first-order conditions for an equilibrium are then that for all $j \neq k \in \{1, 2\}$:

$$\left[1 - r\left(a_j^*\right)\right]\varphi_j(0)\frac{h_\varepsilon}{h_j + h_\varepsilon}W_j + S_j'(a_j^*) - c_j'(a_j^*) = r'\left(a_j^*\right)\left[\left[1 - r\left(a_k^*\right)\right]\phi_j(0) + r\left(a_k^*\right)\right]W_j. \tag{A.7}$$

Consider the special case without injury risk, that is, $r(a) = 0$ for all a, first. Making use of the normality of $\varphi_j(\cdot)$, the first-order condition defining a_j^* can be rewritten as

$$\frac{1}{\sqrt{2\pi}\sigma(h_j, h_k, h_\varepsilon)}\exp\left(-\frac{(m_k - m_j)^2}{2\sigma^2(h_j, h_k, h_\varepsilon)}\right)\frac{h_\varepsilon}{h_j + h_\varepsilon}W_j + S_j'\left(a_j^*\right) - c_j'\left(a_j^*\right) = 0, \tag{A.8}$$

which is equivalent to

$$\frac{1}{\sqrt{2\pi}\sigma(h_j, h_k, h_\varepsilon)}\exp\left(-\frac{|m_k - m_j|^2}{2\sigma^2(h_j, h_k, h_\varepsilon)}\right)\frac{h_\varepsilon}{h_j + h_\varepsilon}W_j + S_j'\left(a_j^*\right) - c_j'\left(a_j^*\right) = 0. \tag{A.9}$$

The latter condition depends on $\Delta \equiv |m_1 - m_2|$ but not on m_1 and m_2 individually. As is apparent from the first-order conditions, $\lim_{\Delta \to -\infty} a_j^* = \lim_{\Delta \to \infty} a_j^* = a_j^n$ when there are no injury risks. Assuming that the second-order condition for a maximum holds,[29] the implicit function theorem implies that

$$sign\left(\frac{da_j^*}{d\Delta}\right) = sign\left(\varphi_j(0)\left(\frac{-2\Delta}{2\sigma^2(h_j, h_k, h_\varepsilon)}\right)\frac{h_\varepsilon}{h_j + h_\varepsilon}W_j\right). \tag{A.10}$$

It follows directly from (A.10) that $\frac{da_j^*}{d\Delta} < 0$ for $\Delta > 0$, and that $\frac{da_j^*}{d\Delta} = 0$ for $\Delta = 0$, in which case j's equilibrium winning probability, $\phi_j(0)$, is equal to $\frac{1}{2}$.[30]

[29] It is easy to check that the second-order condition always holds for small enough Δ in the model without injury concerns.

[30] With more than two contestants, the analysis is considerably more complex. In particular, an agent's effort incentive is no longer maximal if his prior reputation is the same as that of his rivals (assuming the rivals all have the same prior reputations). Rather, the agent's effort incentive will be maximal if he has an advantage over his rivals and an equilibrium winning probability between $\frac{1}{n}$, where n is the number of contestants, and $\frac{1}{2}$.

If $r' > 0$ and the second-order condition holds, then

$$sign\left(\frac{da_j^*}{d(m_j - m_k)}\right) \tag{A.11}$$

$$= sign\left(\left[1 - r\left(a_j^*\right)\right] \frac{d\varphi_j\left(0\right)}{d(m_j - m_k)} \frac{h_\varepsilon}{h_j + h_\varepsilon} - r'\left(a_j^*\right)\left[1 - r\left(a_k^*\right)\right] \underbrace{\frac{d\phi_j\left(0\right)}{d(m_j - m_k)}}_{>0}\right). \tag{A.12}$$

Since the mean of φ_j is $m_k - m_j$, $\frac{d\phi_j(0)}{d(m_k - m_j)} < 0$ which implies $\frac{d\phi_j(0)}{d(m_j - m_k)} > 0$. Overall, the second term in (A.12) is therefore always negative. As implied by the discussion of the situation without injury concerns, $\frac{d\varphi_j(0)}{d(m_j - m_k)} = 0$ for $m_j = m_k$, $\frac{d\varphi_j(0)}{d(m_j - m_k)} < 0$ if $m_j > m_k$, and $\frac{d\varphi_j(0)}{d(m_j - m_k)} > 0$ if $m_j < m_k$. Since the second term in (A.12) is negative, we can conclude that $\frac{da_j^*}{d(m_j - m_k)} < 0$ whenever $m_j \geq m_k$. If j has a (weak) advantage over k, then further improvements in j's relative position reduce j's effort. In the limit where j is certain to win conditional on remaining injury-free, $\lim_{(m_j - m_k) \to \infty} \phi_j\left(0\right) = 1$ and $\lim_{(m_j - m_k) \to \infty} \varphi_j\left(0\right) = 0$, hence the first-order condition in (A.7) directly implies that $\lim_{(m_j - m_k) \to \infty} a_j^* < a_j^n$. By continuity, agents with high enough equilibrium nomination probabilities will exert lower than normal effort as well.

For $m_j < m_k$, on the other hand, the impact of a reduction in asymmetry, i.e., of an increase in $(m_j - m_k)$, is ambiguous. If r is not too steep, then equilibrium effort is increasing in equilibrium winning probability initially but decreasing thereafter. Not also that in the limit case where j has virtually no chance of winning, $\lim_{(m_j - m_k) \to -\infty} \phi_j\left(0\right) = \lim_{(m_j - m_k) \to -\infty} \varphi_j\left(0\right) = 0$, so the first-order condition in (A.7) implies $\lim_{(m_j - m_k) \to -\infty} a_j^* = a_j^n$.

Appendix B: Additional Tables

Table 12: German Euro 2008 team nominations and values of pastselect. **Bold letters indicate nominated players.**

player	pastselect		World Cup 06
	end of 07/08 season	two-year average	
Thomas Hitzlsperger	86.67%	60.24%	Yes
Per Mertesacker	80.00%	59.66%	Yes
Kevin Kuranyi	73.33%	33.43%	-
Clemens Fritz	66.67%	37.94%	-
Marcell Jansen	66.67%	56.97%	Yes
Philipp Lahm	66.67%	75.44%	Yes
Arne Friedrich	60.00%	73.93%	Yes
Bastian Schweinsteiger	60.00%	81.99%	Yes
Lukas Podolski	60.00%	74.25%	Yes
Mario Gomez	60.00%	21.92%	-
Piotr Trochowski	60.00%	34.97%	-
Simon Rolfes	60.00%	22.94%	-
Miroslav Klose	53.33%	71.49%	Yes
Roberto Hilbert	53.33%	21.52%	-
Torsten Frings	46.67%	80.45%	Yes
Bernd Schneider	40.00%	74.42%	Yes
Gonzalo Castro	33.33%	16.00%	-
Manuel Friedrich	26.67%	37.07%	-
Mike Hanke	20.00%	25.68%	Yes
Tim Borowski	20.00%	43.08%	Yes
Christian Pander	13.33%	5.66%	-
Heiko Westermann	13.33%	2.45%	-
Jan Schlaudraff	13.33%	12.82%	-
Alexander Madlung	6.67%	8.67%	-
Jermaine Jones	6.67%	2.67%	-
Paul Freier	6.67%	3.88%	-
Stefan Kiessling	6.67%	4.02%	-
Fabian Ernst	0.00%	0.78%	-
Gerald Asamoah	0.00%	10.58%	Yes
Malik Fathi	0.00%	10.42%	-
Patrick Owomoyela	0.00%	2.50%	-
Sebastian Kehl	0.00%	10.67%	Yes

Notes: The table includes all German players with positive average values of pastselect, except for goalkeepers. No German player without any national team nominations during the sample period was nominated.

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